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Essays in Public Finance and Local Public Policies

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in Economics

by

Wookun Kim

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ABSTRACT OF THE DISSERTATION

Essays in Public Finance and Local Public Policies

by

Wookun Kim

Doctor of Philosophy in Economics

University of California, Los Angeles, 2020

Professor Pablo David Fajgelbaum, Chair

This dissertation includes two essays in applied microeconomics. In Chapter 1, I investigate how much people value local government spending. This is important to measure to better inform government policies and theories of public finance and urban economics, especially when considering how to allocate government spending across locations. To estimate this, I build a quantitative spatial general equilibrium model and combine it with the empirical environment of South Korea where I can leverage a quasi-natural experiment of tax policy reforms to estimate the valuation. I find that an extra dollar of local government spending is valued at 75 cents of their private consumption equivalent. Having obtained the estimate, I embed the measurement into a broader model of the South Korean economy and ask a broader question that involves a general equilibrium analysis of the optimal fiscal transfers across locations: what is the best way to transfer tax revenue across locations in the context where this revenue would be used to finance local government spending. What I find is that fiscal arrangements with small redistribution relative to the actual extent of redistribution observed in South Korea would have positive aggregate effects on welfare. However, completely eliminating the transfer scheme would result in a large welfare loss. In addition to these substantive findings, this chapter has a methodological contribution. The key aspect is to account for two forms of mobility: where

people choose to live, or migration, and where people choose to work, or commuting, which have been thus far studied separately. Throughout my analysis, I show that accounting for both of these margins of mobility is key to correctly estimating the valuation for local government spending and measuring fundamental parameters in the spatial economics literature, which also appear in my framework, namely the elasticities of migration and commuting with respect to spatial frictions. In Chapter 2, I examine the effects of pro-natalist cash transfers on fertility outcomes in South Korea. I exploit the rich cross-sectional variation in cash-transfer generosity over time using 15 years to identify the causal effects of these transfers on the number of births and their health outcomes. Overall, the results provide evidence that cash transfer is an effective policy measure to increase completed fertility and the number of children every born per woman without adversely impacting infant health outcomes and sex composition at birth. Decomposing the birth rates by parity, I find that cash transfers offered for a specific birth parity only affected the parity-specific birth rates. Furthermore, the cash transfers did not change the fertility rate of adolescents.

The dissertation of Wookun Kim is approved.

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2020

I dedicate this dissertation to my family, my love, Kelly, and Eva.

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Chapter 1

The Valuation of Local Government Spending: Gravity Approach and Aggregate Implications

Wookun Kim, UCLA[1]

How much do people value local government spending? What are the effects of fiscal transfers that finance this spending? I develop a spatial equilibrium framework where people's simultaneous migration and commuting choices reveal preferences. I combine this framework with administrative data from South Korea and leverage plausibly exogenous variation in local government spending across districts induced by national tax reforms in 2008 and 2012. The estimated mobility responses imply that workers value each additional dollar of per-capita local government spending by 75 cents of their after-tax income. The general-equilibrium counterfactuals imply that a fiscal arrangement with lower redistribution would result in aggregate gains. A key aspect of my analysis is that bilateral migration and commuting decisions are made jointly. Ignoring any of these margins biases the estimates of preferences for public goods, and of distance elasticities of migration or commuting which play a central role in quantitative spatial models.

[1] Link to most recent version: www.wookunkim.com/research. I am extremely grateful to my advisor Pablo Fajgelbaum for his guidance and support. I thank Adriana Lleras-Muney, Kathleen McGarry, and Jonathan Vogel for their encouragement and suggestions. Youssef Benzarti and Manisha Shah provided valuable advice. I acknowledge Do Young Yoon, Brett McCully, and many other participants at Southern Methodist University, The HKUST, GRIPS, LKYSPP NUS, UCLA applied-micro and international seminars, the UCSB Applied Micro Lunch, the USC-UCI-UCLA Urban Research Symposium, the Warwick Ph.D. Economics Conference, the LACAE, and the All CA Labor Economics Conference at UCSC for helpful comments. This project was supported in part by the UCLA Ziman Center's Rosalinde and Arthur Gilbert Program in Real Estate, Finance and Urban Economics and the California Center for Population Research at UCLA with the grant (T32HD007545; P2CHD041055) from the NICHD. The content is solely my responsibility and does not represent the official views of the NICHD and the NIH. All errors are mine.

1.1 Introduction

How much do people value local government goods? Answering this question is important to inform questions in public finance and urban economics. In this paper, I implement a new gravity approach to measure these preferences. I implement this approach in the context of South Korea, a highly decentralized economy with large heterogeneity in local spending across districts. Then, I compute the optimal levels of fiscal redistribution implied by this measurement.

In the spirit of Tiebout (1956), my approach to estimate the valuation of local government spending is based on mobility. The key novel feature is that preferences for government spending are revealed by people's bilateral migration and commuting decisions in a context where moving is costly. Accounting jointly for both margins of mobility—migration and commuting—is important because these choices are linked and may respond to government spending. Workers may move to places with generous provision of local government goods, but they may also find places attractive to live in if they facilitate access to jobs via commuting. Furthermore, the location of origin (i.e., from where a worker migrates) may influence the choice of both residence and workplace. Using a gravity equation that captures these margins and quasi-natural variation in government spending, I find that the marginal valuation of a dollar of local government spending is equal to 75 cents of disposable income. A key takeaway from my analysis is that ignoring any of these dimensions (place of origin, place of residence, and workplace) biases the estimates of preferences for public goods and of distance elasticities of migration or commuting which play a central role in quantitative spatial models.

The empirical setting of this paper is South Korea. There are three key aspects of the South Korean economy that make it an ideal environment for my analysis. First, local government spending varies across 222 granular spatial units. These spatial units, referred to as *districts*, partition the mainland of South Korea.[2] Each district has a local government which provisions local public goods. This local spending is financed via income tax from its residents, a part of which is locally retained while the rest is redistributed across districts.[3] Second, national tax policy reforms in 2008 and 2012 reduced

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[2] Districts in this paper correspond to 222 administrative units in South Korea called *Si*, *Gun*, or *Gu*. To give a sense of the scale, the total land area of South Korea is about 1 percent of the U.S. or about the same size as the state of Kentucky.

[3] The national government of South Korea redistributes local tax revenue across districts via intergovernmental transfers. The Local Subsidy Act describes a set of formulas computing the amount of intergovernmental transfers each district receives. The rules of redistribution favor districts with lower amenity values and higher population density. There was no major changes in the formula since its enactment in 1994. See Section 1.2.3 for more details.

the income tax rates providing a quasi-natural experiment to estimate the preferences for local government spending. Although these reforms modified national tax policies, they resulted in differential changes in local government revenues because each district had a different socio-economic composition that determined tax base. Due to budget balancing, these changes in local government revenues led to equivalent changes in its spending. Third, I can observe bilateral migration and commuting decisions every 5 years from 2005 to 2015. I use restricted-access administrative data from the Population Census of South Korea to construct a geo-coded panel data set of the number of workers in terms of three locations, which I define by their district of residence 5 years ago, current district of residence, and workplace district.

To guide the analysis, I use a quantitative spatial equilibrium model with a number of features in common with Ahlfeldt et al. (2015) and Monte et al. (2018). As in their frameworks, workers decide where to work and where to live taking wages and floor-space prices into account. The model accommodates an arbitrary number of spatial units (corresponding to the districts in my data). Districts are different in terms of local amenities as a residence and in terms of productivity as a workplace, while commuting costs vary by district pair. The supply of floor space is endogenously determined for commercial or residential use. Following these frameworks, I also incorporate idiosyncratic preferences for residential and employment locations in the spirit of McFadden (1974) and Eaton and Kortum (2002).[4]

In addition to these standard features, I incorporate two key margins. First, as in Morten and Oliveira (2018), workers are heterogeneous in terms of their place of previous residence, which empirically corresponds to where people lived 5 years ago. This margin implies additional spatial frictions on top of commuting. Specifically, I allow the spatial frictions between previous and current residences (i.e., migration) and between previous residence and workplace location (i.e., job finding). Second, residential decisions depend on local government spending, which are financed through a fiscal transfer scheme that corresponds to what I observe in Korea. Local government spending may lead to agglomeration or congestion spillovers depending on the extent of rivalry associated with how much people benefit from local government goods.[5] The framework generates a gravity equation, which expresses the fraction of workers from each origin living in a residence and commuting to a workplace location as a function of spatial frictions between these locations, spatial frictions with respect to the previous residence, wages at the workplace,

[4] See Redding and Rossi-Hansberg (2017) for a review of quantitative spatial models.

[5] E.g., under full rivalry, local government goods are simply publicly provided private goods. Under no rivalry, these goods behave like an agglomeration externality.

and government spending, density and home prices at the residence.

I implement a two-step approach to estimate people's valuation for government goods and spatial frictions. First, I recover reduced-form elasticities governing worker mobility. My identification strategy to estimate the preferences for local government spending is to compare how the decisions of migration and commuting changed due to an increase in local government spending, *ceteris paribus.* I construct instrumental variables based on the tax reforms discussed above and historical residential density to estimate the reduced-form elasticities of worker mobility with respect to local government spending, residential density, and home prices. I find that the probability of migration increases by 1.07 percent for a 1 percent increase in local government expenditure and decreases by 0.8 percent and 0.5 percent for a 1 percent increase in residential density and home prices, respectively. The results imply that there is rivalry associated with local government goods (i.e., a dollar tax in contribution of a worker is shared with his fellow residents to a certain extent). These estimates, nonetheless, are not enough to recover the valuation of local government spending. For that, I also need to know how much people move in response to spatial frictions and wages. Using the mobility response to wages, I can then express the response to government spending in disposable-wage equivalent units and recover and estimate the utility parameters.

In the second step, I use the structural properties of the model and estimate the effects of spatial frictions (i.e., costs of migration, commuting, and job finding) and wages on worker mobility. I estimate the distance elasticities of migration, commuting, and job finding to be negative and stable over time. I show that estimating the distance elasticity of migration while not taking commuting into account and the distance elasticity of commuting without accounting for migration lead to large biases. With respect to the distance elasticity of migration, the upward bias arises because workers migrate over long distance when they are better compensated from the local labor market at the destination.[6] Estimating the distance elasticity of commuting while not accounting for migration leads to an overestimation because there are additional costs associated with commuting due to the costs of migration and job finding in addition to the direct cost of commuting explained by the distance of commuting.[7] These biases primarily arise

[6] For example, Bryan and Morten (2018) estimates the distance elasticity of migration based on the migration patterns in the U.S. and Indonesia without taking commuting into account. Based on the migration pattern in South Korea alone, I estimate a value of the elasticity similar to theirs, which is about 4.7 times smaller in magnitude than my estimate based on both migration and commuting patterns.

[7] Ahlfeldt et al. (2015) estimates this elasticity in the context of the City of Berlin, Germany. My estimate following their estimation strategy is similar to wht they found, which is 2.1 times larger in magnitude (more negative) than the estimate based on both migration and commuting.

because the previous literature studied migration and commuting under the assumption that previous residential location does not affect where workers work, which the data unequivocally rejects.

Following the approach in Ahlfeldt et al. (2015), I estimate how much people move in response to wages (i.e., the Fréchet shape parameter). Using this estimate, I re-scale the estimated reduced-form elasticity of worker mobility with respect of local government spending and compute the marginal valuation of local government goods in dollars. As a result, I find that workers on average value an additional dollar of local government goods equal to 75 cents of their after-tax income. This estimate is similar to the point estimate of Suárez-Serrato and Wingender (2014), who use a different source of variation in the U.S. context.[8] In order to correctly estimate how much people value local government goods, it is important to take both margins of mobility into account especially when spatial units are finely defined. The valuation of local government spending is biased upward if migration is not taken into account and downward if commuting is not taken into account.

Using the estimated general equilibrium model, I quantify the welfare consequences of the spatial distribution of local government spending. I conduct a set of counterfactual policy experiments to shed light on the optimal modes of fiscal decentralization (local taxation vs. redistribution). Across counterfactual exercises, I vary the extent of redistribution, i.e., how much local government spending depends on redistributive intergovernmental transfers relative to local taxation. Many countries around the world (e.g., Canada, Germany, Australia and Japan) make fiscal transfers across regions, similar to the South Korean system featured in this paper. I allow for counterfactual regimes to mimic what is observed in other countries, ranging from a high redistribution (as in Canada and Denmark) and little redistribution (as in the U.S.).

I find that there would be a welfare improvement if the extent of redistribution observed in 2015 were reduced. However, the complete elimination of the redistributive intergovernmental transfers would lead to a sizable loss of welfare. The results indicate that transfers of income are too high from fiscally strong districts (i.e., districts with higher average income) to the weak (i.e., districts with lower average income) under the redistribution policy observed in 2015. The benefit of the transfers in the net-receiving districts is dominated by the loss in the net-contributing districts. Lastly, I show that different assumptions on spatial mobility of workers (e.g., costless migration and pro-

[8] Suárez-Serrato and Wingender (2014) estimate the valuation of government spending in the context of the U.S. They exploit the population revisions following the decennial Census ("Census Shock") and the measurement error in population levels during non-Census years to isolate exogenous variation in federal spending across counties in the U.S.

hibitively costly commuting) call for a significantly different extent of redistribution. For example, if no spatial frictions of migration and job finding are assumed as in the commuting literature, a fiscal arrangement with significantly lower redistribution appears optimal. In this scenario, a lower extent of redistribution improves the overall welfare by reducing the incentive for workers to reside in net-receiving districts at the expense of longer commute.

My paper builds upon several existing literatures. The public finance literature examines the effects of government policies on the spatial distribution of workers. Tax differentials across space incentivize workers to move across the state and country borders (Kleven et al., 2014; Akcigit et al., 2016; Moretti and Wilson, 2017). Some papers estimate positive amenity values for government spending and regulations from housing prices (Cellini et al., 2010; Black, 1999; Chay and Greenstone, 2005) in the spirit of Rosen (1979) and Roback (1982). There are only a few papers that directly estimate how much workers value government goods using government spending. Using a spatial general equilibrium framework, Suárez-Serrato and Wingender (2014) estimate the effect of federal spending on local economies in the U.S. by exploiting changes in population levels used to determine the size of federal funding for localities due to Census shocks (Suárez-Serrato and Wingender, 2016). Fajgelbaum et al. (2019) rely on tax differences across U.S. states over time and the spatial proximity to other states to estimate worker preferences for government expenditure.[9]

My approach includes various novel features relative to these papers. First, the spatial unit used in this paper is finer than the spatial units commonly considered in the literature (e.g., states and county groups in the U.S.). Given the granular spatial units, I leverage both migration and commuting patterns to estimate how much workers value local government goods and services. Second, I provide a new identification strategy using national tax reforms as a source of plausibly exogenous variation in local government spending to estimate the elasticity of worker mobility. Third, I estimate the effect of residential density on worker mobility by following the standard approach used in the urban economics literature to estimate agglomeration and congestion forces (Ciccone and Hall, 1996; Combes and Gobillon 2015; de la Roca and Puga, 2017).

This paper also contributes to the fiscal decentralization literature. The majority of the papers in this literature focus on theoretically and empirically examining the consequences of the changes in fiscal autonomy of local governing entities.[10] There are relatively

[9] Gelbach (2004) focuses on the female population in the U.S. eligible for state welfare programs and finds that the interstate migration patterns of this population are not sensitive to the distribution of welfare benefits across states.

[10] For instance, Fisman and Gatti (2002) documents that fiscal decentralization leads to a lower level of

few empirical papers studying the effects of policy instruments employed for fiscal decentralization (e.g., local taxation and redistribution). Government goods and services are often public, thus creating fiscal spillovers. Wildasin (1980) finds that households may locate in an optimal fashion in the presence of the spatial distribution of local government spending and notes that fiscal spillovers may result in non-optimality. Fajgelbaum and Gaubert (2018) characterize the optimal transfers for efficient allocations and the policies implementing the transfers. Albouy (2012) presents a theoretical framework to determine efficient and equitable transfers across localities and evaluates the welfare consequences of the equalization policy in Canada. I contribute to the literature on fiscal decentralization by computing the optimal mix of location taxation and redistribution.

Lastly, this paper contributes to a growing literature on quantitative economic geography models. There are a number of recent papers that have studied the migration and commuting decision, separately. In the case of migration, Bryan and Morten (2018) study the cost of migration as a source of friction that results in labor market misallocation using the case of Indonesia. Morten and Oliveira (2018) quantify the impact of transport networks using the construction of a radial highway system in Brazil when workers can migrate across space. Moretti and Wilson (2017) estimate the negative effects of tax rate differences across states on the migration of star-scientists in the U.S.[11] In the case of commuting, Ahlfeldt et al. (2015) and Tsivanidis (2019) study the commuting patterns and their contributions to the spatial distribution of economic activity in the city of Berlin, Germany and in the city of Bogotá, Columbia, respectively. The literature on migration assumes that workers live and work in the same locations. The literature on commuting assumes often implicitly zero spatial frictions associated with migration and job finding. Monte et al. (2018) have a notion of migration in addition to commuting; however, they assume that where workers migrate from does not affect their commuting decisions. To the best of my knowledge, this paper is the first to present a spatial equilibrium model featuring both bilateral migration and commuting in the economic geography literature. Furthermore, these papers concerning the geographical mobility of workers do not study the roles of the public sector.

The remainder of this paper is structured as follows. In Section 1.2, I describe the data sources and the key aspects of the South Korean economy. In Section 1.3, I present a

corruption. Bianchi et al. (2019) show that fiscal decentralization led to a higher female labor force participation because local governing authorities expanded nursery schools. See Oates (1999) for a broader literature review on fiscal federalism.

[11] There are more papers studying the migration patterns in the U.S., Vietnam, and Brazil based on spatial equilibrium models: e.g., Piyapromdee (2017); Albert and Monras (2019); Balboni (2019); Pellegrina and Sotelo (2019). My model abstracts away from the dynamic model presented in Caliendo et al. (2019).

partial equilibrium model in which workers choose where to live and where to work in the presence of local government goods and services as well as costs associated with mobility. Then, I estimate the elasticities of worker mobility with respect to local government spending, residential density, and home prices in Section 1.4. Section 1.5 focuses on the effects of spatial frictions on worker mobility and estimate three reduced-form elasticities measuring the responsiveness of worker mobility (migration, commuting, and job finding) with respect to distance between localities. In Section 1.6 and 1.7, I embed the partial equilibrium model presented in Section 1.3 into a general equilibrium setup and describe how I parameterize the model. Finally, I consider counterfactual policy experiments concerning the extent of redistribution and its aggregate welfare implications in Section 1.8. Section 1.9 concludes.

1.2 Data and Background

In this section, I discuss some key aspects of the South Korean economy and the data I have collected to study how the spatial distribution of local government spending affects the spatial mobility of workers and, more broadly, how it affects the aggregate welfare of workers. Specifically, in Section 1.2.1, I discuss main data sources of the key variables for my empirical study. In Section 1.2.2, I define the geographic units used in this study and document the migration and commuting patterns in South Korea. Lastly, Section 1.2.3 discusses the national policies on local public finance and describe national tax reforms in 2008 and 2012.

1.2.1 Data

The observed spatial distribution of workers is a consequence of decisions on two margins of geographical mobility of workers—migration and commuting. Therefore, my empirical analysis has a specific data requirement. First, I need data that records worker's previous residence, current residence, and current employment location. Second, the data has to be spatially representative. Third, local government spending should vary at the same spatial disaggregation across which workers activley make both migration and commuting decisions. South Korea is one of the few countries, which meet all of the requirements.

My main data source for the spatial distribution of workers is the restricted Population Census of South Korea. The Population Census of South Korea is conducted every five years and sample 20 percent of the entire population. I use the three most recent waves from 2005, 2010, and 2015. I restrict the sample to working male household heads between the ages of 25 to 60, who commute a round trip of less than 180 kilometers.[12] The sample size is about 3.5 million households. The Census questionnaire asks district of residency five year ago, current district of residence, and district of workplace location. Based on this information, I construct a panel data set of the distribution of workers by residence five years ago, current residence, and workplace location.

Data on local government spending was collected from the administrative data (Yearbook of Local Public finance) from the Ministry of Interior and Safety of South Korea. I digitized local government information by total revenue and revenue from the following sources: local income taxes and intergovernmental transfers. This information allows me to recover the share of the intergovernmental transfer each locality received from the national government in a given year. In addition, the Ministry of Land, Infrastructure, and Transport publishes the land price fluctuation rates for each district. I collected this information for 2005, 2010, and 2015. The fluctuation rate is defined as a ratio of the average land price in a given year to the average land price in 2004.

I supplement the main data set with local characteristics in 2015 using the administrative data from various government agencies to complete the parameterization of the spatial general equilibrium model I present later in the paper. The two key variables are wages and housing prices.[13] A major limitation of the Population Census is that the information on wealth and income is not surveyed. Instead, I use the Economic Census of 2015, which surveys the universe of establishments, and compute the average annual wages in each district. The Ministry of Land, Infrastructure, and Transport maintains the universe of housing transactions from 2006 to 2015. I construct district-level prices per unit of floor space in 2015 by employing a Case-Shiller type repeated sales approach at the district level, similarly done in Ahlfeldt et al. (2015). Lastly, I compute distance between every pair of districts by connecting their centroids.[14]

[12] The reason for restricting the same to male household heads is motivated by the fact that migration decisions are made at the household level. Over 90 percent of the households in the Population Census have male household heads. The female labor force participation in South Korea is one of the lowest among the OECD countries (Lee, 2017). The age restriction is to only include workers who have completed education. Also, less than 1 percent of workers report a commuting distance over 90 kilometers in each direction.

[13] See Appendix 1.11 for the complete list of additional variables and their sources.

[14] In addition to the key variables explained above, I collect other local characteristics (e.g., land use, suicide rates, divorce rates, and number of firms) for cross-validation exercises and over-identification checks carried out later in the paper. See Appendix 1.11 for details.

1.2.2 Spatial Mobility in South Korea

South Korea has the 11th largest economy in the world, comparable to Canada and Spain in terms of GDP and GDP per capita, respectively. While South Korea is only about 1 percent of the U.S. geographically, the population level was as high as 51 million in 2015, about 16 percent of the population in the U.S.

The spatial units used in this paper are districts in South Korea, which are the smallest administrative units with local governing authority. I will hereon refer to these district-level governing entities as local governments, I focus on the 222 contiguous districts that partition the South Korean mainland, excluding the districts of Jeju Island.[15] The average size of each district is 224,310 in terms of population (91,471 households), approximately twice as large as the average population of a county in the U.S.

I describe two dimensions of spatial mobility—commuting and migration—in South Korea. First, I begin with the commuting patterns in South Korea. Workers in South Korea spend about 7.3 percent of their workday commuting between their residence and workplace locations, reflecting the commuting patterns documented in Monte et al. (2018) for the U.S. and Schafer (2000) for 26 countries around the world.[16]

In Panel A of Table 1.1, I report summary statistics on commuting patterns in 2005, 2010, and 2015. On average, about 27 percent of residents work outside their district of residence and about 30 percent of workers commute to work from other districts. In addition, I plot the fraction of residents commuting to other districts against distance between residence and workplace in Panel (a) of Figure 1.1. I observe that the probabilities of commuting decreases in distance. This implies that the cost associated with commuting increases in distance consistent with prior literature (Ahlfeldt et al., 2015; Monte et al., 2018; Tsivanidis, 2019).

Second, with respect to migration, about one in seven households migrate across district borders annually; the implied annual inter-district migration rate is around 13 percent.[17] Aggregated at the province level (i.e., 16 groups of districts), the annual migration rate in South Korea is 5 percent, similar to the inter-county migration rate in the U.S. reported

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[15] As of 2005, there are 226 local governments, four of which were split as a consequence of redistricting. I keep the administrative units consistent to the administrative boundaries set in 2005 by defining groups of districts for those which underwent redistricting.

[16] See Redding and Turner (2015) for further discussion on cost of commuting and transportation costs.

[17] I do not observe annual migration patterns in the Population Census. Instead, the annual migration rates are calculated using the restricted-use administrative records of the universe of migrants in South Korea during the same time period (2005-2015). The migrant records are not used for analysis in this paper because it does not provide information on where migrants work.

in Molloy et al. (2011).[18] In Panel B of Table 1.1, I report the probabilities of migration over 5 years at the district level. On average, about 19 percent of residents in a district are migrants who have migrated from other districts within 5 years, while 18 percent of residents have migrated out of their residence in the past 5 years. Panel (b) of Figure 1.1 plots the probabilities of migration conditional on location of origin against the distance between the origin and current residence. In line with the literature on migration, I also observe that the probability of migration decreases with distance (Bryan and Morten, 2018; Morten and Oliveira, 2018).

In Figure 1.2, I plot the number of households in the figure on the left and local expenditure on the right by district. There are districts with generous local expenditure and many households. This pattern suggests that workers are more likely to reside in districts with more generous provision of local government goods and services. Additionally, local determinants like wages, home prices, and amenities also influence worker's migration and commuting decisions. The spatial distributions of these additional forces also explains the distribution of workers across localities. Figure 1.3 provide suggestive evidence that workers are willing to migrate further and commute longer to live in a district with a relatively higher level of local government expenditure and lower home prices.

1.2.3 Local Government Revenue and Tax Reforms in 2008 and 2012

The total local government expenditure accounts about 8 percent of South Korean GDP in 2015. I focus on two main sources of local government revenue: local income taxes and intergovernmental transfers, which constitute 14 percent and 72 percent of local government spending, respectively.[19] Panel C of Table 1.1 reports the summary statistics of local government expenditure for 222 districts in the years 2005, 2010, and 2015. The average total local expenditure is 363 billion KRW (approximately 363 million USD); the average per-capita local expenditure is 7,638 USD, widely ranging from 906 USD to 29,622 USD.[20]

[18] Therefore, the contrast in annual migration rates between South Korea (inter-district) and the U.S. (inter-county) can be explained by geographical size differences between U.S. counties and South Korean districts.

[19] The remaining 14 percent of local government revenue is comprised of non-tax receipts (e.g., fees, charges, and fines) and borrowing, last of which is only about 0.06 percent on average. Hereon, I refer to the sum of local income taxes and intergovernmental transfers as local government revenue or expenditure.

[20] For simplicity, I will continue assuming the unit of government spending (and wages) in USD throughout the remainder of the paper.

There is substantial spatial variability in the degree to which districts depend on local income taxes and intergovernmental transfers for their total local spending. Local income tax constitutes on average about 16.8 percent of the total local revenue. Intergovernmental transfers constitutes about 83.2 percent of total local spending.[21] The share of local government revenue from local income tax ranges from 2.1 percent to 56 percent.

The national fiscal policies (progressive income tax system and extents of fiscal decentralization and redistribution) and local tax bases (number of workers and their income) determine local government revenue. The Local Autonomy Act—first enacted in 1949—was revived in 1991 after 30 years of suspension due to military dictatorships that ended in 1987. The purpose of the Act was to "strive for democracy and efficiency of local autonomous administration and to ensure balanced development of local areas..." (Local Autonomy Act, 1991). The national government amended the Local Tax Act and Local Subsidy Act to enable local autonomy in 1994. The Local Tax Act and Local Subsidy Act together with the Income Tax Act promulgates progressive income tax rates. I will refer to the collection of these three Acts as the national fiscal policies.

The national fiscal policies determine the size of local governments in two ways. First, local governments collect income tax from their residents according to the income tax rates outlined in the Income Tax Act, which is uniform across all districts. Local governments retain a fixed share of their income taxes and deliver the rest to the national government. I refer to the fixed share as *local-national revenue sharing* and the amount of income tax revenue left at the local level as *local income tax*. Lastly, the national government allocates a fixed share of its tax revenue for redistribution. Then, the national government makes *intergovernmental transfers* to each local government, calculated by a set of formula determining the shares of the total fund allotted to each local government.[22] In sum, the extent of fiscal decentralization (the fraction of total tax revenue local governments spend) and the rules of redistribution (how to allocate intergovernmental transfers across

[21] In Appendix 1.12.1, Figure 1.9 plots the spatial distribution of local tax revenue in Panel (a) and the ratios of local tax revenue to total local spending in Panel (b). Likewise, Panel (c) and (d) plots the spatial distribution of intergovernmental transfers and the contribution of intergovernmental transfers to total local spending for each district. According to Figure 1.9, there is a considerable variation in how fiscally strong localities are.

[22] The Local Subsidy Act details the formula employed to determine how much intergovernmental transfers to be rebated to each locality. The overarching objective of intergovernmental transfers is to help develop "the public administration of local governments in a sound manner with the adjustment of their finances by subsidizing financial resources necessary for the public administration of local governments" (Local Subsidy Act, 1994). There are a number of countries both developed and developing (e.g., Germany, UK, Canada, Australia, and India) with a similar local finance instrument (equalization grants) to promote balanced financial capacities horizontally. While the U.S. does not have a federal system directly aiming to reduce differences in fiscal capacities across localities, many of the federal grants and policies have features that are implicitly equalizing across states and localities (e.g., EITC, SNAP, medicare, and medicaid).

local governments) determine local government revenues.

Tax Reforms

there were two major reforms on national tax policies: one in 2008 and the other in 2012. Figure 1.4 plots the marginal income tax rates before and after the tax reforms in 2008 and 2012. In 2008, the Income Tax Act was amended to substantially decrease income tax rates across income brackets: from 11 percent to 8.8 percent for the low income bracket (annual income less than 12 million KRW); from 22 percent to 18.7 percent for the middle income group (12 million to 46 million KRW); and from 33 percent to 28.2 percent for the high income group (46 million to 88 million KRW).[23] In 2012, the national government further reduced the income tax rates to 6.6 percent for the low income group, to 16.5 percent for the middle, and to 26.4 percent for the high group. The tax reforms did not affect the rules of redistribution outlined in the Local Subsidy Act. Figure 1.10 in Appendix 1.12.1 plots the current shares of intergovernmental transfers to the shares 5 years ago. The estimated slopes comparing the redistribution policies in a given year to these five years ago are close to 1.

1.3 Discrete Choice Model of Worker Location Decisions

In this section, I present a discrete choice model, in which workers make decisions on migration and commuting. In the model, a worker decides where to live and where to work, taking wages, prices of residential floor space, local government goods and services, and the location choices of all other workers into account. My model is different from the spatial equilibrium models commonly used in the recent literature examining the spatial mobility of workers in two ways. First, similar to Fajgelbaum et al. (2019), I augment the model by introducing goods and services provisioned by local governments. Second, the model features both commuting (Ahlfeldt et al., 2015; Tsivanidis, 2019) and migration decisions (Bryan and Morten, 2018; Morten and Oliveira, 2018) which have been independently studied. There are iceberg costs of worker mobility rising from three spatial frictions: migration, commuting, and job finding. The key prediction of the

[23] The total number of income brackets had been four until the second amendment in 2012, which introduced one additional income brackets for the even richer. For my analysis, I focus on the lowest three income brackets which include more than 95 percent of workers in South Korea according to the Ministry of Strategy and Finance of South Korea. I also note that the first reform in 2008 resulted in small changes in the cutoffs of each bracket to account for inflation since the last change in 1994. Since the first reform, the cutoffs for the lowest three income brackets remain the same.

model is a gravity equation which summarizes the distribution of workers in terms of initial residence, current residence, and workplace location.

1.3.1 Model Environment

The whole economy has R measure of workers (also, interchangeably referred to as residents) and comprises of J discrete number of spatial units (i.e., districts), indexed by r for current residence, m for workplace, and o for initial residence. As a residence, each district is characterized by exogenous local amenities B_r, per-unit floor space price Q_r, and local government goods and services g_r. As a workplace, a district is characterized by wage w_m, which is subject to income tax. Therefore, workers commuting to district m receive after-tax income equal to $(1 - \tau_m)w_m$ for their private consumption of the single final good c_{rm} and residential floor space h_{rm} in residence of district r. In addition, there are iceberg costs of worker mobility across space in three dimensions: migration D_{or}, commuting D_{rm}, and job finding D_{om}, summarized in a single disutility index $D_{orm} = \varepsilon_{orm} D_{or} D_{rm} D_{om}$ where ε_{orm} is a stochastic error term following a log-normal distribution with its mean equal to 1. The first two spatial frictions follow the standard formulations in the literature on migration and commuting. The spatial friction captures the iceberg cost associated with finding a job in workplace location m from initial residence o.

Workers are born in or assigned to initial residence o (also, interchangeably referred to as origin). The initial distribution of workers is given by π_o. I assume each worker inelastically supplies one unit of labor. After observing an idiosyncratic utility shock for every possible pair of residence r and workplace m, a worker chooses a residence-workplace pair that maximizes his utility given his initial residence, after-tax wages, floor space prices, local government expenditure, location choices of other workers, local amenities, and the iceberg costs of mobility.

1.3.2 Worker's Location Decisions

The preferences of a worker i are defined over amenities, consumption of the single final good, consumption of floor space for housing, public goods, and the iceberg costs associated with migration, commuting, and job finding. The direct utility of a worker i who chooses to move from his initial residence o to a new residence r and commutes to a workplace m is $z_{irm} u_{orm}(c_{rm}, h_{rm})$, where c_{rm} is consumption of the single final good (numeraire) and h_{rm} is consumption of floor space for housing.

First, $u_{orm}(c_{rm}, h_{rm})$ corresponds to the systemic component of the preference and follows the Cobb-Douglas form:

$$u_{orm}(c_{rm}, h_{rm}) = \frac{B_r}{D_{orm}} \left(\frac{c_{rm}}{\beta}\right)^{\beta} \left(\frac{h_{rm}}{1-\beta}\right)^{1-\beta} g_r^{\lambda}. \tag{1.1}$$

Amenity fundamental B_r captures intrinsic residential characteristics that make district r more or less attractive to live in (e.g., the weather, beaches, and scenic views). The parameter β determines the share of expenditure on the final consumption good.[24] Following Fajgelbaum et al. (2019) and Albouy (2012), real government expenditure enjoyed by each worker living in district r g_r is local government expenditure G_r, normalized by a function of the total number of workers living in district r R_r:

$$g_r = \frac{G_r}{R_r^{\theta}}. \tag{1.2}$$

The parameter θ controls the extent to which local government goods and services are rival and ranges from 0 if non-rival (pure public good) to 1 if rival (publicly-provided private good). The parameter $\lambda \geq 0$ captures the weight of local government goods and services in preferences relative to consumption of the single final good and floor space for housing.

Given unit price of floor space for housing Q_r and after-tax wage $(1-\tau_m)w_m$, the budget constraint is $c_{rm} + Q_r h_{rm} = (1-\tau_m)w_m$. The Cobb-Douglas preference implies that β share of after-tax wage is allocated to the consumption of the single final good and the rest to the consumption of residential floor space. Therefore, the indirect utility of a worker i choosing to live in district r and commute to district m is $V_{iorm} = z_{irm} v_{orm}$,where

$$v_{orm} = \frac{B_r(1-\tau_m)w_m}{D_{orm} Q_r^{1-\beta}} \left(\frac{G_r}{R_r^{\theta}}\right)^{\lambda}. \tag{1.3}$$

The systemic component of the indirect utility increases in local amenities B_r, wage w_m, and local government expenditure G_r, while it decreases in per-unit price of floor space Q_r, residential population R_r, and spatial frictions $D_{orm} = \varepsilon_{orm} D_{or} D_{rm} D_{om}$. The composite iceberg cost associated with migration, commuting, and job finding D_{orm} enters the indirect utility function multiplicatively. Therefore, there is an isomorphic formulation in which after-tax wages are reduced due to commuting and job finding costs, and amenity values are decreased due to migration cost.

Second, z_{irm} is an idiosyncratic preference shock that captures the idea that each individual worker has idiosyncratic reasons to find a residence and a workplace more or less attractive. I model this heterogeneity in preference in spirit of McFadden (1974) and

[24] Davis and Ortalo-Magné provide empirical evidence supporting the constant housing expenditure share, using the U.S. as a case study.

Eaton and Kortum (2002). For a worker i residing in district r and working in workplace location m, the idiosyncratic component of his utility is drawn from an independent Fréchet distribution:

$$\Pr(z_{irm} < z) = \exp(-T_r M_m z^{\epsilon}), \tag{1.4}$$

where the parameter $T_r > 0$ determines the average utility of living in district r; the parameter $M_m > 0$ determines the average utility of working in district m; and the shape parameter $\epsilon > 1$ governs the dispersion of the utility draw.[25] Then, the distribution of workers living in district r and working in district m by initial residence o:

$$\pi_{orm} = \frac{\left(\frac{\tilde{B}_r(1-\tau_m)\tilde{w}_m}{D_{orm}Q_r^{1-\beta}}\left(\frac{G_r}{R_r^{\theta}}\right)^{\lambda}\right)^{\epsilon}\pi_o}{\sum_{r'=1}^{J}\sum_{m'=1}^{J}\left(\frac{\tilde{B}_{r'}(1-\tau_{m'})\tilde{w}_{m'}}{D_{or'm'}Q_{r'}^{1-\beta}}\left(\frac{G_{r'}}{R_{r'}^{\theta}}\right)^{\lambda}\right)^{\epsilon}} \equiv \frac{\Phi_{orm}\pi_o}{\Phi_o}, \text{ where } \Phi_o = \sum_{r'=1}^{J}\sum_{m'=1}^{J}\Phi_{orm}. \tag{1.5}$$

Because some of the unobserved local characteristics (i.e., T_r and M_m) always appear in the gravity equation together with unobserved local amenities B_r and wages w_m, I define the following composite terms denoted by a tilde: adjusted amenities $\tilde{B}_r = B_r T_r^{1/\epsilon}$ and adjusted wages $\tilde{w}_m = w_m M_m^{1/\epsilon}$.

Workers are more likely to live in a residences with a high amenity value and local government expenditure and lower per-unit floor space price, net of congestion/agglomeration forces and migration costs.[26] Workers are more likely to commute to workplace locations with higher after-tax wages net of commuting and job finding costs.

1.4 Key Reduced-Form Elasticities of Worker Mobility

In this section, I estimate the reduced-form elasticities of worker mobility with respect to local government expenditure, residential density, and floor space prices derived from

[25] The indirect utility V_{irmr_0} is Fréchet distributed since V_{iorm} is a monotonic function of the Fréchet distributed idiosyncratic preference shock z_{irm}. The maximum utility is itself Fréchet distributed appealing to the stability postulate.

[26] In Appendix 1.14.1, I discuss how I derive the gravity equation (1.5). It is also general enough to produce the gravity equations summarizing the spatial distribution of workers that the literature on commuting and migration have considered based on economic geography models (Ahlfeldt et al., 2015; Bryan and Morten, 2018; Morten and Oliveira, 2018; Monte et al., 2018; Moretti and Wilson, 2017). In Appendix 1.14.2, I show that the gravity equations used elsewhere can be derived based on the gravity equation (1.5).

the observed distribution of worker mobility in South Korea. Section 1.4.1 discusses an econometric specification, which I derive using the gravity equation, a key prediction of the spatial equilibrium model presented in the preceding section. In order to consistently estimate the reduced form elasticities of interest, I exploit the episodes of national tax reforms discussed in Section 1.2.3 as well as information on the historical residential density. In Section 1.4.2 and 1.4.3, I present the estimation results and discuss the interpretation and robustness of the estimated reduced-form elasticities.

1.4.1 Estimation Strategy

The gravity equation (1.5) describes how workers sort across districts in terms of residential and workplace locations from previous residence. I take the log transformation of both sides of the gravity equation and obtain the following econometric specification by augmenting the terms with time subscript whenever applicable to permit the panel structure of the data:

$$\ln \pi_{orm,t} = \phi_{om,t} + \phi_{or} + \phi_{rm} + \underbrace{\lambda\epsilon}_{\beta_G} \ln G_{r,t} - \underbrace{\theta\lambda\epsilon}_{\beta_R} \ln R_{r,t} - \underbrace{(1-\beta)\epsilon}_{\beta_Q} \ln Q_{r,t} + \zeta_{orm,t}. \quad (1.6)$$

The coefficients in front of log local government expenditure ($\beta_G = \lambda\epsilon$), log number of workers living in r ($\beta_R = \theta\lambda\epsilon$), and log prices of floor space ($\beta_Q = (1-\beta)\epsilon$) are the reduced-form elasticities in question and are functions of structural parameters. The job finding fixed effects interacted with year dummy variables $\phi_{om,t}$ flexibly capture the workplace-specific factors (e.g., after-tax wages and average utility from working in district m) and the factors specific to the origins (e.g., number of workers who used to live in o and the denominator of the gravity equation (1.5)) as well as the iceberg cost of job finding. The migration fixed effects ϕ_{or} and the commuting fixed effects ϕ_{rm} capture the time-invariant component of the iceberg costs of migration and commuting as well as the intrinsic residential characteristics of district r that makes it a more or less attractive place to live in.[27] Lastly, the error term $\zeta_{orm,t}$ includes the rest of the factors in equation 1.5 (i.e., adjusted amenities and time-varying stochastic components of the spatial frictions net of costs associated with job finding).

The errors in Equation 1.6 can be correlated in two ways. First, there is a classic clustering concern explained in Moulton (1990). Second, one may worry about the serial

[27] Note that land area for each district is absorbed into the migration and commuting fixed effects because area is a time-variant feature of each locality. This implies that β_R can be interpreted as the elasticity of worker's mobility with respect to residential density.

correlation over time within a panel dimension Bertrand et al. (2004). In order to address these concerns, I report standard errors that are robust to heteroskedasticity and allow multi-way clusterings. I allow errors to correlate across previous residences and across workplace locations sharing the same current residence in a given year. In addition, the serial correlation within each of the panel dimension (a triplet of previous residence, current residence, and workplace location) over time.

1.4.1.1 Fixed Effects

The mapping between the econometric specification (1.6) and the gravity equation from the spatial model (1.5) helps to understand potential confounders and consequent biases. First, the job finding fixed effects interacted with year dummy variables $\phi_{om,t} = \ln(1-\tau_{m,t})^{\epsilon}\tilde{w}_{m,t}^{\epsilon}\exp(-\delta\epsilon d_{om})\pi_{o,t}/\Phi_{o,t}$ control for the benefit from choosing to work in m net of the job finding cost from previous residence o. Workers are more likely to choose workplaces with higher net benefits. Given higher returns from workplace location, workers are willing to accept a lower amount of local government spending at their residential location. Furthermore, worker's valuation of a given workplace location depends on their origin because for example they rely on their network to find higher paying jobs, and this network is usually formed at the origin (Card, 2001; Cadena and Kovak, 2016). Thus, if one does not control for the different levels of the attractiveness of the nearby workplace by origin, the OLS estimate of β_G will be downward biased.

Next, a higher net labor market return attracts residents. This positive correlation between the residential density and the labor market return biases the OLS estimate of β_R upward. Workers with higher after-tax wages would be able to afford higher housing prices. Similarly, excluding the job finding by year fixed effects biases the OLS estimates of β_R and β_Q because residential density and home prices partially reflect the fact that there are attractive workplaces nearby for workers of a given origin.

Second, omitting the migration fixed effects $\phi_{or} = -\rho\epsilon d_{or}$ and commuting fixed effects $\phi_{rm} = -\kappa\epsilon d_{rm}$ are likely to bias the OLS estimate of β_G downward. While the costs of migration and commuting inhibit worker mobility, workers may choose residences with higher local government expenditures to offset their migration and commuting costs. Districts that are attractive to live in are likely to have higher housing prices and residential densities. If so, the costs of migration and commuting are likely correlated positively with the residential density and housing prices, again in the sense of compensating differentials. Then, OLS estimates of β_R and β_Q would be biased downward.[28]

[28] Note that local government expenditure, residential density, and housing prices are correlated with each other. It is useful to have a sense of the potential directions of bias in the conventional way

1.4.1.2 Endogeneity

Even after conditioning on the set of fixed effects discussed above, OLS estimates of β_G, β_R, and β_Q suffer from endogeneity due to omitted variable bias and measurement errors. The error term $\zeta_{orm,t} = \ln \tilde{B}^{\epsilon}_{r,t}\varepsilon^{-\epsilon}_{orm,t}$ includes the adjusted local amenity values. Local government expenditures and local amenities are likely negatively correlated because redistributive intergovernmental transfers favor places with low amenity values *ceteris paribus*. This negative correlation between amenity values and government expenditures would generate a downward bias in the OLS estimate of β_G. Next, districts with higher amenities attract inflows of migrants, which lead to a higher residential population. This means the OLS estimate of β_R would be overestimated. Lastly, high amenity values would be priced into home prices in the sense of hedonic pricing. Then, the OLS estimate of β_Q would suffer from an upward bias in this case toward zero.

Furthermore, there is an additional concern of measurement error with respect to Q_r. I do not directly observe the home prices for 2005, 2010, and 2015. Instead, I use data on land prices as a proxy. Assuming classical measurement error, an OLS estimate of coefficient β_Q would be attenuated. In fact, because all of the endogenous regressors are correlated with each other, all the other OLS estimates would also be biased.

1.4.1.3 Instrumental Variables

Because the estimating equation (1.6) has three endogenous variables, in order to consistently estimate their coefficients, I propose three instrumental variables based on the national tax reforms and the historical values of residential density. For each district r, I first construct two instrumental variables, exploiting the episodes of tax reforms in 2008 and 2012 discussed in Section 1.2.3:

$$IV^b_{r,t} = \tau_{b,t}\pi_{b|r,2000}, \tag{1.7}$$

where the income tax rates $\tau_{b,t}$ change over time; subscript b denote each of the two income brackets I use (low and high). The values of $\tau_{b,t}$ are unique in each year t and income bracket b because the reforms took place between the years when the Population Census was conducted (2005, 2010, and 2015). Furthermore, I leverage the variation in the predetermined share of workers by educational attainment level b, symmetrically defined in

by thinking about the relationship between omitted variables and the dependent variable and the relationship between omitted variables and the endogenous regressors. Nevertheless, the covariances among the endogenous variables as well as their relationship with an omitted variable need to be taken into account in order to properly characterize the directions of potential omitted variable bias. The system of equations summarized in (1.38) in Appendix shows the complexity of the omitted variable bias with three endogenous variables.

terms of two levels (low for workers who have completed high school at most and high for workers with some college degrees), for each district in 2000 $\pi_{b|r,2000}$. The predetermined local educational distribution proxies the distribution of workers by income brackets in each district.[29]

The instrumental variables ($IV_{r,t}^{low}$ and $IV_{r,t}^{high}$) capture the tax contributions of low and high income groups predicted by income distribution in 2000. Therefore, by construction, the relevance of the instrumental variables follows immediately from the local government budgetary structure: government expenditures increase in tax contributions. To satisfy the exclusion restriction, the instrumental variables must not directly influence workers to prefer one residence over another, except through their impacts on local government expenditures, floor space prices, and residential densities. There are two sources of variation in the proposed instruments. One source is tax rate changes over time. Conditional on wages ($\phi_{om,t}$), workers are subject to the same tax rates regardless of their residential and employment locations. Thus, the tax rates do not directly affect their location decisions. Another source is the cross-sectional variation in the educational distribution within each district in 2000. Although my model does not take a stance on the sorting by skill levels, the previous literature has found that workers sort based on education or skill levels as skill-mix determines residential amenities (Eeckhout et al., 2014; Diamond, 2016; Fajgelbaum and Gaubert, 2018). The validity of the proposed instruments still holds as long as the tax reforms changed the educational composition in each district. In Appendix 1.12.6, I confirm that the tax reforms are orthogonal to changes in educational composition within each district over time. Therefore, the instrumental variables constructed based on the national fiscal policy reforms remain valid with potential sorting by skill levels.

Second, the last instrumental variable $IV_{r,t}^{R}$ is based on the historical residential density as previously used in Ciccone and Hall (1996) and de la Roca and Puga (2017).. Specifically, I use the natural logarithm of the number of households in r thirty years ago ($\ln R_{r,t-30}$) as the data allows a lag of up to 30 years: the number of households in 1975, 1980, and 1985. As Combes and Gobillon (2015) explain, historical values of residential density are usually considered relevant due to inertia in local population as local housing stock and infrastructure last over time. They are also believed to be exogenous to contemporaneous local characeteristics that affect worker mobility because the changes in the type of economic activity and historical events like war reshape the

[29] The sum of the proposed instrumental variables ($\sum_b IV_{r,t}^{b}$) shares the same structure as Bartik-type instruments widely used on the literature Bartik (1991). Borusyak et al. (2018) and Goldsmith-Pinkham et al. (2018) discuss sources of variation in shift-share instruments and identification approaches. See Adão et al. (forthcoming) for inference procedures when employing shift-share instruments.

economic landscape.[30] The validity of the instrument hinges on the assumption that historical residential densities do not directly affect the worker location decisions today.[31] This assumption is violated in the unlikely situation in which workers rely on the population levels 30 years ago, instead of its contemporaneous or more recent levels, when deciding where to live today.

To consistently estimate the reduced-form elasticities of worker mobility with respect to local government expenditure, residential density, and home prices (β_G, β_R and β_Q), I use the two-stage least squares (2SLS) estimator with the following identification assumption:

$$E\begin{bmatrix} IV_{r,t}^{low}\zeta_{orm,t} & |\phi_{om,t},\, \phi_{or},\, \phi_{rm} \\ IV_{r,t}^{high}\zeta_{orm,t} & |\phi_{om,t},\, \phi_{or},\, \phi_{rm} \\ IV_{r,t}^{R}\zeta_{orm,t} & |\phi_{om,t},\, \phi_{or},\, \phi_{rm} \end{bmatrix} = 0. \tag{1.8}$$

1.4.2 Estimation Results

In Table 1.2, I report the OLS estimates of the elasticities of mobility to local government expenditure, residential density, and home prices. In Column (1), I report the OLS estimates without including any fixed effects. The OLS estimate of β_G is negative, against the expectation that workers value local government goods. The estimated coefficient in front of the log number of households is 0.12, which implies strong agglomeration. According to the estimated coefficient of β_Q, a 1 percent increase in home prices decreases worker mobility by 0.042 percent.

In Column (2), I report the OLS estimates with the fixed effects of job finding interacted with year dummy variables. Compared to the estimate in Column (1), the OLS estimate of β_G increases to 0.097. This increase can be explained by netting out the negative correlation between local government expenditures and labor market returns from redistributive intergovernmental transfers. Furthermore, the estimated elasticity of worker mobility with respect to residential density decreases to 0.061, implying that there is a positive association between after-tax wages discounted by the cost associated with job finding and residential density in line with intuition. Lastly, the OLS estimate of β_Q increases to -0.01; however, this estimate is statistically not different from zero. On the one hand, the increase in the estimate of β_Q is against the direction of bias associated with omitting the fixed effects. On the other hand, the estimated value is likely a result of an attenuation bias due to measurement error.

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[30] In the case of South Korea, a series of military dictatorship lasted about three decades until 1987.

[31] The validity of the historical residential density as an instrumental variable can be also justified using the demographic balancing equation used in demography (Preston et al., 2000).

In Column (3) and (4), I gradually add the fixed effects of migration pairs and commuting pairs to purge out the confounding effects of costs associated with migration and commuting on worker mobility. Because compensating differentials imply a positive correlation between the costs of mobility and local government expenditures, the coefficient estimate of β_G should increase as a result of the additional fixed effects. However, the OLS estimate of β_G changes little in Column (3) and (4). The result reflects the omitted variable bias towards zero from unobserved local amenity values, which are negatively correlated with local government expenditures and positively affects worker mobility. With respect to the estimates of β_R and β_Q, the OLS estimates increase compared to the estimated values reported in Column (2) in line with the potential directions of bias discussed.[32]

Table 1.3 summarizes the two-stage least squares results in Column (2), (3), and (4) and compares the results with the OLS results in Column (1). Note that the OLS estimates reported in Column (1) are the same as the ones in Column (4) of Table 1.2. First, according to the estimates in Column (2), log local government expenditure is positively correlated with the predicted tax contributions from the low and high income groups, IV^{low} and IV^{high}. The magnitudes of the estimates are similar because both tax contributions are measured in KRW. The lag residential density IV^R is also positively correlated with log local government expenditure. Second, the current residential density is positively correlated with the predicted tax contributions, but negatively correlated with the historical residential density. Conditional on the set of fixed effects, a negative coefficient in front of the historical residential density implies that the districts that grew at higher rates 30 years ago currently grow relatively slower. The last first stage result concerns the home prices. Log home prices are positively correlated with the predicted tax contributions as well as the historical residential density. All the the coefficients reported in Column (2), (3), and (4) are statistically different from zero at the 1 percent level. To formally test the strength of the first stage results, I compute various F-stats including SW conditional F-stats, which test the explanatory power of the excluded instruments in the presence of multiple endogenous variables (Sanderson and Windmeijer, 2016; Stock et al., 2002). I report the SW conditional F-stats and verify the strength of the first stages.

[32] The estimates in Column (4) is based on the fully saturated specification (1.6). According the estimated coefficients in Column (4), worker mobility increases by 0.1 percent with respect to 1 percent increase in local government expenditure and by 0.59 percent with respect to 1 percent increase in residential density. The estimated elasticity of worker mobility with respect to home prices is not only statistically insignificant, but also economically small. The OLS estimates are contaminated by measurement errors in home prices and the omitted variable bias from excluding local amenity values that make residence more attractive and are correlated with the included regressors.

In Column (5) of Table 1.3, I report the 2SLS estimates of the elasticities of worker's mobility to local government expenditure, residential density, and home prices. First, the estimated elasticity to local government expenditure is statistically different from zero and substantially larger compared to the OLS estimate in Column (1). As discussed earlier, this large increase implies that there is a substantial downward bias rising from omitting time-varying local amenities, which are negatively correlated with local government expenditures, but make residences more attractive. The result indicates that one percent increase in local government expenditure increases the probability of worker's mobility (equivalently the conditional probability of migration) by 1.07 percent.

Second, the estimated elasticity of worker mobility with respect to residential density becomes negative. However, I cannot reject the null hypothesis that the estimate is different from zero. Statistical insignificance notwithstanding, the change in the sign of the elasticity indicates that there is a considerable bias toward zero resulting from omitting local amenities which the 2SLS strategy addresses. According to the estimate, a 1 percent increase in residential density leads to a 0.844 percent decrease of the conditional probability of migration.

Based on the structural relationship between the estimated elasticities, I further estimate the value of the structural parameter $\theta = -\beta_R/\beta_G$, which capture the extent of rivalry associated with local government goods and services, by the Delta method. The estimated value of parameter θ is 0.787 with a standard error equal to 0.315. The magnitude of the estimate suggests a non-negligible effect of rivalry from residential density.

Lastly, the estimated elasticity of worker mobility with respect to home prices is equal to -0.49, substatially larger than the OLS estimate in Column (1). As explained earlier, there are two sources of bias to the OLS estimate of β_Q. One is the omitted variable bias. Because amenity values and home prices are positively correlated, the OLS estimate in Column (1) is biased upward towards zero. The other is measurement errors, attenuating the effect of home prices on worker mobility towards zero. The 2SLS estimates which correct for these issues show that the conditional probability of migration decreases by 0.49 percent as home prices increase by 1 percent.[33]

1.4.3 Interpretation of Estimates

The elasticity of worker's mobility to government expenditure has not been extensively estimated in the previous literature. Suárez-Serrato and Wingender (2014) estimate 1.46

[33] In Table 1.13 in Appendix, I report the estimation results based on an alternative, parsimonious specification in which time-varying origin-workplace fixed effects are replaced with fixed effects separately for time-by-origin, time-by-workplace, and origin-pair. The results are robust.

for the elasticity of population at the county group level by leveraging exogenous variation in federal spending in the U.S. Fajgelbaum et al. (2019) obtains a similar value of the elasticity based on the number of workers at the state level in the U.S. Although the comparison is not perfect since they consider different source of variation, time periods, and geography, my estimate of 1.07 is close to their estimates.[34]

The literature on agglomeration economies includes population density as part of amenities and productivity (Ciccone and Hall, 1996; Glaeser and Mare, 2001; Ahlfeldt et al., 2015; de la Roca and Puga, 2017). The magnitude of its effect has been estimated to be positive, but rather small; the existing values of agglomeration parameter range from 0.01 to 0.06. This being said, the congestion parameter via local government goods θ in my model includes the agglomeration force.[35] Overall, I find that local government goods and services are rival. However, since it is not fully rival, a tax contribution from an additional resident is shared with all the other residents. Therefore, the effect of residential density on worker mobility is net agglomerating.

Lastly, the estimation results suggest that spatial frictions reflected in the iceberg costs of migration, commuting, and job finding are important determinants of worker's location decisions. In Table 1.11 in Appendix, I report the OLS estimates including each of the fixed effects separately and show that the effects of local government expenditures, residential density, and home prices on worker mobility is sensitive to job finding costs in Column (2), migration costs in Column (3), and commuting costs in Column (4).

In Table 1.14 in Appendix, I compare the OLS and 2SLS estimates based on both migration and commuting flows in Column (1) and (2) to the OLS and 2SLS estimates based on migration flows alone in Column (3) and (4) and commuting flows alone in Column (5) and (6). The 2SLS estimates in Column (4) and Column (6) are biased because the exclusion restriction in each case is violated. On the one hand, based on the migration pattern alone, the effect of local government spending on the probability of migration is underestimated because workers move to places with higher commuting potentials, which compensate the lack of local government spending. On the other hand, the same effect is overestimated by about 5 times based on the commuting pattern alone because, in addition to direct cost of commuting, there exist migration and job finding costs that enable each commute. This set of estimation results emphasizes the importance

[34] My estimate of β_G is slightly smaller than the ones estimated in Suárez-Serrato and Wingender (2014) and Fajgelbaum et al. (2019). This is likely because they study the effects of government expenditures that affect firms as well as workers.

[35] There is a simple isomorphic formulation in which the agglomeration force is directly featured in the model. If the local amenities B_r is endogenous and depends on amenity fundamentals b_r and residential density R_r^γ, where γ captures the residential agglomeration force. Then, the reduced-form parameter $\beta_R = (\theta\lambda - \gamma)\epsilon$ and $\beta_R/\beta_G = \theta - \gamma/\lambda$.

of jointly accounting for both migration and commuting and for proper conditioning to consistently estimate the elasticities of worker mobility with respect to local government spending, residential density, and floor space prices.

1.5 Estimation of Spatial Frictions

Spatial frictions make it difficult for workers to reallocate across space. The model presented in Section 1.3 features the iceberg costs of worker mobility including three spatial frictions: the costs associated with migration, commuting, and job finding. In this section, I estimate the effects of spatial frictions on the spatial mobility of workers. I shed light on the importance of jointly considering migration and commuting decisions in correctly estimating the distance-elasticities of migration and commuting.[36]

1.5.1 Spatial Frictions in Migration and Commuting Decisions

I rewrite the gravity equation (1.5) by grouping the location-specific factors by residence ϕ_r, by workplace location ϕ_m, and by previous residence ϕ_o:

$$\pi_{orm} = \frac{\phi_o \phi_r \phi_m}{\underbrace{(\varepsilon_{orm} D_{or} D_{rm} D_{om})^{\epsilon}}_{D_{orm}}}. \tag{1.9}$$

I refer to Equation (1.9) as a generalized gravity equation of migration and commuting as this equation generalizes the gravity equations in the literature on migration and commuting.

Based on Equation (1.9), the expression for the spatial distribution of workers by their origins and current residences is given by:

$$\pi_{or} = \frac{\phi_o \phi_r}{D_{or}^{\epsilon}} \underbrace{\sum_{m=1}^{J} \frac{\phi_m}{(\varepsilon_{orm} D_{rm} D_{om})^{\epsilon}}}_{ALMA_{or}\varepsilon_{or}}. \tag{1.10}$$

The key difference between the expression above (1.10) and the one considered in the literature on migration is the last term $\sum_{m=1}^{J} \phi_m/(\varepsilon_{orm} D_{rm} D_{om})^{\epsilon}$. This additional term can be expressed in terms of stochastic ε_{or} and systemic components. I refer the systemic component of the additional term to as *augmented labor market access* ($ALMA$). $ALMA$ shares a similar structure with the labor market access (LMA) in Morten and Oliveira

36 Recovering unobserved factors in the model requires the estimates of these elasticities as discussed in Section 1.7.3.

(2018) and more generally with the market access approach in Donaldson and Hornbeck (2016), but includes an additional factor D_{om}. On the one hand, the conventional LMA has a unique value for each of current residences (i.e., destinations) since it captures the benefit of accessing the local labor market net of commuting costs. On the other hand, $ALMA$ allows LMA to vary by previous residences (i.e., origins) to account for heterogeneous costs of job finding and captures the benefit of accessing the local labor market net of both commuting and job finding costs.

$ALMA$ captures the idea that workers from different origins value the same local labor market of a residence differently due to the cost of job finding. The extent to which workers can benefit from the labor market of a certain residence may depend on where they migrate from due to, for instance, a migrant network that makes job finding easier for workers from a certain origin relative to those from somewhere else (Card, 2001; Cadena and Kovak, 2016). Although $ALMA$ does not explicitly appear in the gravity equations used in the migration literature, $ALMA$ provides an important information about how workers sort across space. Workers conditional on their origins are more likely to migrate to a residence with higher $ALMA$, while a higher value of $ALMA$ enables workers to afford a higher cost of migration.

Second, the literature on commuting employs a gravity equation, which summarizes the spatial distribution of workers in terms of their current residential and workplace locations. By summing π_{orm} in Equation (1.9) over initial residences, I obtain a gravity equation that characterizes the commuting patterns of workers:

$$\pi_{rm} = \frac{\phi_r \phi_m}{D_{rm}^{\epsilon}} \underbrace{\sum_{o=1}^{J} \frac{\phi_o}{(\varepsilon_{orm} D_{or} D_{om})^{\epsilon}}}_{AMMA_{rm}\varepsilon_{rm}}. \tag{1.11}$$

Again, the key difference between the gravity equation above (1.11) and the one considered in the literature on commuting is the last term $\sum_{o=1}^{J} \phi_o/(\varepsilon_{orm} D_{or} D_{om})^{\epsilon}$. This term can be written in terms of stochastic ε_{rm} and systemic components, last of which I term *augmented migrant (worker) market access* ($AMMA$). $AMMA$ captures the average appeal of a commute (between a residence and a workplace location) for migrants net of costs associated with migration and job finding. Therefore, there are two types of costs that explain the commuting patterns One type is a usual direct cost of commuting D_{rm}. The other is an indirect cost that captures the idea that it is costly to move to residence r and find a job in workplace location m from previous residence o in order to commute between r and m. Similar to the direct cost of commuting, this indirect cost makes the appeal of a commute less attractive.

$AMMA$ measures how accessible each commute is for workers originating from different places on average and varies at the commute-pair level.[37] On the one hand, it is likely to see more workers carrying out a certain commute when this commute has a higher value of $AMMA$. On the other hand, if the commute is costly, the appeal of this commute is lower and so is $AMMA$. While the literature on commuting is silent about the role of $AMMA$ as a determinant of commuting decisions, accounting for $AMMA$ is important to correctly estimate the distance elasticity of commuting.

1.5.2 Estimation Strategies and Results

I take a step towards evaluating how much spatial frictions quantitatively explain the spatial distribution of workers observed from the Population Census of South Korea. As defined in Section 1.3, I impose a structure on each of the bilateral linkages such that these linkages depend on distances d_{jk} between localities j and k as similarly done in, for instance, Morten and Oliveira (2018) for the cost of migration and Ahlfeldt et al. (2015) for the cost of commuting:

$$D_{or} = \exp(\rho d_{or}),\ D_{rm} = \exp(\kappa d_{rm}),\ D_{om} = \exp(\delta d_{om}). \tag{1.12}$$

The parameters ρ, κ, and δ control the sizes of migration, commuting, and job finding costs with respect to distances between spatial units. The motivation for imposing the same structure on the cost of job finding as the costs of migration and commuting is that finding a job is harder for workers who are located farther away from potential job sites. Taking into account that the data is available for cross-sections of 3 years (2005, 2010, and 2015), I augment the gravity equation by adding time subscripts:

$$\pi_{orm,t} = \frac{\phi_{r,t}\phi_{m,t}\phi_{o,t}}{\varepsilon^{\epsilon}_{orm,t}\exp(\rho\epsilon d_{or} + \kappa\epsilon d_{rm} + \delta\epsilon d_{om})}. \tag{1.13}$$

I estimate the reduced-form elasticities of worker mobility with respect to distances ($\rho\epsilon$, $\kappa\epsilon$, $\delta\epsilon$) using the South Korean Census. [38]

1.5.2.1 Cost of Migration with respect to Distance

I take the log transformation of both sides of the generalized gravity equation (1.13)with time subscripts and the structure of bilateral linkages to obtain the expression as follows:

[37] The term *augmented migrant market access* reflects a concept that different commutes have differential capacities to access and attract migrants.

[38] In Appendix 1.12.2, I conduct a type of decomposition exercise to shed light on the contribution of the spatial linkages jointly and discretely to the observed variation in the spatial distribution of workers. The spatial linkages individually explain about 23 to 70 percent of the observed variation.

$$\ln \pi_{orm,t} = \phi_{rm,t} + \phi_{om,t} - \rho\epsilon d_{or} + \varepsilon^{mig}_{orm,t}, \tag{1.14}$$

where the current residence by workplace fixed effects interacted with year dummies $\phi_{rm,t}$ capture time-varying location specific factors at the current residence $\ln \phi_{r,t}$, the workplace $\ln \phi_{m,t}$, and the cost of commuting $-\kappa\epsilon d_{rm}$. The origin by workplace fixed effects interacted with year dummies $\phi_{om,t}$ capture time varying location specific factors at origin $\ln \phi_{o,t}$ as well as the cost of job finding $-\delta\epsilon d_{or}$. The parameter $\rho\epsilon$ is the semi-elasticity of migration flows with respect to distances of migration. The expected sign of $-\rho\epsilon$ is negative because workers are less likely to migrate to places that are farther away. The last term $\varepsilon^{mig}_{orm,t}$ corresponds to the log of the stochastic error $\varepsilon_{orm,t}$; I assume this error term is orthogonal to distances of migration.[39] I allow the errors to be correlated across migration pairs.

In Table 1.4, I start with a simple OLS estimation without any fixed effects and gradually add two sets of fixed effects (commuting pairs $\phi_{rm,t}$ and job finding pairs $\phi_{om,t}$), one at a time. The estimate in Column (1) without any fixed effects is -0.002, statistically different from zero. This estimate is likely biased from omitting the determinants of migration that are correlated with distance of migration. For instance, if workers migrate longer distances to find better jobs (higher wages), the estimate is biased toward zero. In order to purge out the net benefits of living in r and commuting to workplace m, I include pairwise fixed effects for commuting pairs $\phi_{rm,t}$. The estimated coefficient is now slightly more negative at -0.004, reported in Column (2).

In Column (3), I flexibly control for the cost of job finding by adding pairwise fixed effect for job finding pairs $\phi_{om,t}$; this specification corresponds to Equation (1.14). The estimated semi-elasticity is -0.033 and means that the probability of migration decreases by 3.3 percent with respect to a one-kilometer increase in the distance of migration. The large difference between the estimates in Column (2) and Column (3) implies that there exists a substantial upward bias rising from failing to account for the difficulty in finding jobs for workers who are migration from more distant places. The estimate in Column (3) captures the positive relationship between distance and the cost of migration, net of the costs associated with commuting and job finding.

Given that the distance of migration is a time-invariant feature that links the spatial

[39] I estimate Equation (1.14) using a linear fixed effects estimator. The identification assumption is that, the distances of migration are uncorrelated with all other determinants of residential location choices conditional on the fixed effects. The error term may capture random measurement error in distances of migration. Although I do not observe exact distances of migration, the magnitude of potential measurement errors with respect to distance of migration are likely to be small because the geographical units are defined more finely compared to the spatial units considered in the previous literature.

units, I test whether or not the semi-elasticity of migration to distance varies over time. I include two additional regressors to Equation (1.14): distance interacted with dummy variables for year 2005 and 2010. The coefficients in front of the additional regressors tell us how different the semi-elasticities are in 2005 and 2010 relative to in 2015. The estimation result is reported in Column (4). The magnitudes of the estimated coefficients are economically small and statistically not different from zero. I conclude that the semi-elasticity of migration to distance is relatively constant, and therefore is a time-invariant feature describing the data.[40]

Lastly, I examine the consequence of using the probability of migration, a dependent variable commonly used in the previous literature on migration, to estimate the semi-elasticity of migration with respect to distance. I estimate a specification analogous to what the literature uses to estimate as follows:

$$\ln \pi_{or,t} = \tilde{\phi}_{r,t} + \tilde{\phi}_{ot} - \rho\epsilon d_{or} + \varepsilon^{mig}_{or,t}, \tag{1.15}$$

where the current residence and the origin fixed effects interacted with year dummies ($\tilde{\phi}_{r,t}$ and $\tilde{\phi}_{o,t}$) capture any push and pull factors specific to the origin and current residence that affect migration. To consistently estimate the semi-elasticity of migration to distance $-\rho\epsilon$, the error term $\varepsilon^{mig}_{or,t}$ must be orthogonal to either distance d_{or} or the dependent variable $\ln \pi_{or,t}$, or both. The gravity equation helps to unpack the error term. Based on Equation (1.10) with time subscripts on all the terms except distances, $\varepsilon^{mig}_{or,t}$ corresponds to $\ln ALMA_{or,t}\varepsilon_{or,t} = \ln \sum_{m=1}^{J} \frac{\phi_{m,t}}{(\varepsilon_{orm,t} D_{rm} D_{om})^{\epsilon}}$. An estimate without controlling for the effects of $ALMA_{or,t}$ on migration flows would be biased towards zero because, as explained above, $ALMA_{or,t}$ is correlated positively with both distance and the observed migration flows. I estimate Equation (1.14) and report the estimated coefficient in front of distance in Column (5). Conforming to the expected direction of the omitted variable bias, the estimate is only about a fifth of the estimate in Column (3) because workers are willing to migrate longer distances when they face higher returns from the local labor market at the destination.[41]

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[40] The results are robust to estimating the distance elasticity of migration pooling observations for each year (2005, 2010, and 2015).

[41] The estimate in Column (5) of Table 1.4 falls within the range of available estimates in the literature. Bryan and Morten (2018) estimates the elasticity of migration to distance in the U.S. (-0.553) and Indonesia (-0.717). Re-scaling the estimated semi-elasticity of -0.007 by the average migration distance (75.34 kilometers), the implied elasticity of migration to distance based on my estimate is -0.53.

1.5.2.2 Cost of Commuting with respect to Distance

To estimate the semi-elasticity of commuting with respect to distance, I derive the following specification based on the generalized gravity equation:

$$\ln \pi_{orm,t} = \phi_{or,t} + \phi_{om,t} - \kappa\epsilon d_{rm} + \varepsilon^{com}_{orm,t}, \tag{1.16}$$

where the origin by current residence fixed effects interacted with year dummies $\phi_{or,t}$ capture time-varying location specific factors at the origin $\ln \phi_{o,t}$ and the current residence $\ln \phi_{r,t}$ as well as the cost of migration $-\rho\epsilon d_{or}$; the origin by workplace fixed effects interacted with year dummies $\phi_{om,t}$ capture time-varying location specific factors at the workplace $\ln \phi_{m,t}$ as well as the cost of job finding $-\delta\epsilon d_{om}$. The parameter $-\kappa\epsilon$ is the semi-elasticity of commuting flows with respect to distance of commuting. Because workers are less likely to commute longer distances from their location of residence, the sign of the semi-elasticity must be negative. The stochastic error term $\varepsilon^{mig}_{orm,t}$, orthogonal to distances of commuting, includes the log of the stochastic error. I allow the errors to be correlated across commuting pairs.

Table 1.5 report the estimation results. Like before, I start with a simple OLS estimation without any fixed effects and gradually add two sets of fixed effects (migration pairs $\phi_{or,t}$ and job finding pairs $\phi_{om,t}$), one at a time. The estimate without any fixed effects is -0.013, statistically different from zero in Column (1). This estimate is likely biased from omitting determinants of commuting flows that are correlated with distance of commuting. For example, workers who migrated from places farther away may not want to bear higher commuting costs in addition to cost of migration. Then, the estimate is biased toward zero. In order to account for the omitted variable bias associated with migration cost, I introduce the migration pair fixed effects in Column (2). As expected, the estimate reported in Column (2) is -0.035, more negative compared to the estimate in Column (1). Furthermore, the returns from working in m net of job finding cost, captured by $D_{om,t}$ are positively correlated with the commuting flows and allows workers to afford higher commuting cost. This implies another bias toward zero.

In order to address this issue, Column (3) estimates the semi-elasticity of commuting flows with respect to distance with both fixed effects of migration and job finding pairs. The estimated elasticity in Column (3) is -0.045: a one-kilometer increase in commuting distance decreases the probability of commuting by 4.5 percent.[42] To understand how

[42] Travel time is also widely used to define a cost of geographical mobility (Ahlfeldt et al., 2015; Morten and Oliveira, 2018). In Appendix 1.12.4, I show that travel time associated with commuting has a one-to-one relationship with distance of commuting. I re-estimate Equation (1.16) by using commute time reported in the Population census as an endogenous regressor, instrumented with distance of

stable the semi-elasticity of commuting to distance over time is, I additionally include distance interacted with year dummy variables for 2005 and 2010. The estimation results in Column (4) indicate that the semi-elasticity of commuting with respect to distance is stable over time.

Next, I examine what happens if the bilateral linkages of migration and commuting are not accounted for when estimating the semi-elasticity of commuting with respect to distance. To do so, I follow the literature on commuting and use the probability of commuting $\ln \pi_{rm,t}$ as a dependent variable and estimate the following specification:

$$\ln \pi_{rm,t} = \tilde{\phi}_{r,t} + \tilde{\phi}_{m,t} - \kappa\epsilon d_{rm} + \varepsilon^{com}_{rm,t}, \tag{1.17}$$

where the residence and the workplace fixed effects interacted with year dummies ($\tilde{\phi}_{r,t}$ and $\tilde{\phi}_{m,t}$) capture any factors specific to residence and workplace that affect commuting (costs of living and wages). In order to consistently estimate the semi-elasticity of commuting to distance $\kappa\epsilon$, the error term $\varepsilon^{com}_{rm,t}$ must be uncorrelated to either distance d_{rm} or the probability of commuting $\ln \pi_{rm,t}$, or both. Similar to the case of migration, the log of Equation (1.11) with time subscripts whenever applicable has a direct correspondence with Equation (1.17). The residual term $\varepsilon^{com}_{rm,t}$ is equal to $\ln AMMA_{rm,t}\varepsilon_{rm,t} = \ln \sum_{o=1}^{J} \frac{\phi_{o,t}}{(\varepsilon_{orm,t} D_{or} D_{om})^{\epsilon}}$. It is clear that an increase in $\ln AMMA_{rm,t}$ increases the probability of commuting. Estimating $\kappa\epsilon$ without controlling for the effects of $\ln AMMA_{rm,t}$ on commuting flows would be biased away from zero if a high commuting cost is associated with a low value of $AMMA_{rm,t}$. Column (5) reports the estimated semi-elasticity based on Equation (1.17). The estimate is -0.074, which is more negative compared to the estimate in Column (3) in line with the intuition.[43]

1.5.2.3 Cost of Job Finding with respect to Distance

In this subsection, I estimate the semi-elasticity of job finding with respect to distance. I derive an estimating equation by taking the log transformation of Equation (1.13):

$$\ln \pi_{orm,t} = \phi_{rm,t} + \phi_{or,t} - \delta\epsilon d_{om} + \varepsilon^{jf}_{orm,t}, \tag{1.18}$$

commuting. The estimated semi-elasticity (-0.036) is stastistically not different from the estimate in Column (3) of Table 1.5 (-0.033).

[43] The estimate in Column (5) of Table 1.5 is close to the available estimates in the literature. Ahlfeldt et al. (2015) estimates the same semi-elasticity based the inter-district commuting flows in Berlin, Germany in 2008 contemporaneous to the time period considered in this paper. Their estimated semi-elasticity of commuting with respect to distance (also measured in kilometer) is equal to -0.07.

where the commute-pair fixed effects $\phi_{rm,t}$ capture net benefits of living in r and working in m, $\ln \frac{\phi_{r,t}\phi_{m,t}}{D_{rm}^{\epsilon}}$ (e.g., housing prices, wages, and commuting cost); the migration-pair fixed effects $\phi_{or,t}$ capture the cost of migration $-\rho\epsilon d_{or}$ as well as any factors that make o a more or less attractive residence to stay $\ln \phi_{o,t}$; the sign of the parameter $\delta\epsilon$ is likely positive because it is harder to find jobs that are farther away from where workers migrate; the last term $\varepsilon_{orm,t}^{jf}$ captures the random noise. I allow the errors to be correlated across job finding pairs.

The simple OLS estimate of the semi-elasticity of job finding with respect to without any fixed effects is -0.001, reported in Column (1) of Table 1.6. Workers are willing to accept a high cost of job finding (equivalently, a large d_{om}) if doing so allows them to find a pair of residence and workplace locations with higher wage, lower cost of living and lower commuting costs. These correlations results in bias towards zero. In Column (2), I introduce the pairs fixed effects for residence and workplace locations and find an estimate more negative, compared to Column (1). Furthermore, because the distance of migration and the distance of job finding are positively correlated and workers are less like to migrate farther away, the estimate in Column (2) is still biased toward zero.

Column (3) reports the estimated semi-elasticity of job finding with respect to distance with both fixed effects as prescribed in Equation (1.18). According to Column (3), the probability of job finding decreases by 1.6 percent for a one-kilometer increase in distance of job finding. To examine how stable the semi-elastisticity of job finding is with respect to distance, I introduce distances interacted with year dummy variables for 2005 and 2010. In Column (4), the coefficient estimate for distance reported in the first row is the semi-elasticity of job finding to distance in 2015. The difference between the estimate in Column (4) and the estimate reported in Column (3) is economically small and is not statistically significant.[44]

1.5.3 Implications

Taking the gravity framework to the spatial distribution of workers in South Korea, I find that the spatial linkages between localities (costs of migration, commuting, and job finding) are important determinants of the spatial distribution of workers. In particular,

[44] To the best of my knowledge, there is no existing estimate of the decay parameter $\delta\epsilon$ (i.e. elasticity of job finding with respect to distance) in the literature. That being said, my estimate of the spatial decay of job finding can be considered as a reduced-form parameter combining the effects of distance on job match (employment) and job application (intent for employment), last of which Manning and Petrongolo (2017) estimate based on a spatial model of job search using the data on the demand and supply of the job search process in the U.K. They find a relatively strong decay of job applications in distance. Marinescu and Rathelot (2018) also finds similar results (implied semi-elasticity of job application to distance equal to 0.02) in the context of the U.S.

they are systematically explained by distances between spatial units. The estimated reduced-form elasticities are negative and stable over time.[45]

I make two important distinctions from the previous literature on migration and commuting. First, residence does not need to be a place for both living and working. Second, where workers come from matters for not only determining where they live, but also where they work today.[46] I find substantial biases with the estimates of the distance elasticities of migration and commuting reported in the previous literature.

First, the estimated elasticities of migration available in the literature are likely biased toward zero because the cost of migration is positively correlated with the benefits from changing residences discounted by costs associated with commuting and job finding ($ALMA$). This means that workers appear to be willing to migrate longer distance for better labor market access. Second, the available estimates for distance elasticity of commuting in the literature are likely biased away from zero (more negative). Because of omitting the appeal of commuting net of indirect costs rising from migration and job finding that enable a certain commute ($AMMA$), workers appear to be more sensitive to commuting distance than they actually are.[47]

1.6 Quantitative Spatial General Equilibrium Model

I take a step towards quantifying the welfare consequences of the fiscal arrangements observed in 2015. Accordingly, I embed the partial equilibrium model of worker's location

[45] This finding, in particular related to migration, is consistent with the assumption on migration friction in Caliendo et al. (2019), which build the sectoral mobility costs in Dix-Carneiro (2014). Ahlfeldt et al. (2015) make the same assumption about the semi-elasicity of commuting and applies the same spatial decay of commuting estimated based on the commuting patterns of workers in the city of Berlin in 2008 to explain the commuting patterns before and after the division and reunification of East and West Germany.

[46] Also, Pellegrina and Sotelo (2019) find that the origins of agricultural workers matter in determining the types of crops they cultivate they migrate to a different region in the context of Brazil.

[47] The results also shed light on timings of mobility decisions. Intuitively, there are two alternative timings of how workers decide where to live and where to work. First, a worker may decide a residence where he would like to live (including the option to stay), and then find a job. If this timing is true, the semi-elasticity of job finding should be estimated to zero controlling for the commuting-pair fixed effects. Second, a worker may find a job first, then decide where to commute from. If this alternative timing is true, then the semi-elasticity of commuting should be estimated similarly with or without the fixed effects accounting for the job finding cost conditional on the migration pair fixed effects. Both of these alternative timings are inconsistent with the observed spatial distribution of workers. The findings altogether imply that a certain timing assumption is too restrictive to explain the variations in the observed spatial distribution of workers. Consistent with these findings, the model presented in this paper allows workers to make migration and commuting decisions jointly.

decisions presented in Section 1.3 into a general equilibrium setup. I model the production of consumption goods and the allocation of floor spaces for residential and commercial use. Local government spending is determined based on national policies on taxation, revenue sharing, and the rules of redistribution. In equilibrium, wages, floor space prices, and local government expenditures are endogenously determined along with the spatial distribution of workers. Lastly, I define the spatial general equilibrium of the economy.

1.6.1 More on Worker's Location Decisions

I characterize the market clearing conditions for migration and commuting based on the gravity equation (1.5) derived in Section 1.3. First, summing the probabilities of choosing residence r and workplace m conditional on moving from origin o across workplaces, I obtain the expression for the probabilities of moving to r given origin o:

$$\begin{aligned}\pi_{r|o} &= \sum_{m=1}^{J} \frac{\pi_{orm}}{\pi_o} = \frac{\sum_{m=1}^{J} \Phi_{orm}}{\Phi_o} \\ &= \frac{T_r \left(\frac{B_r}{D_{or} Q_r^{1-\beta}} \left(\frac{G_r}{R_r^\theta} \right)^\lambda \right)^\epsilon \sum_{m=1}^{J} M_m \left(\frac{(1-\tau_m) w_m}{D_{om} D_{rm}} \right)^\epsilon}{\sum_{r'=1}^{J} T_{r'} \left(\frac{B_{r'}}{D_{or'} Q_{r'}^{1-\beta}} \left(\frac{G_{r'}}{R_{r'}^\theta} \right)^\lambda \right)^\epsilon \underbrace{\sum_{m'=1}^{J} M_{m'} \left(\frac{(1-\tau_{m'}) w_{m'}}{D_{r'm'} D_{om}} \right)^\epsilon}_{ALMA_{or'}}}.\end{aligned}$$

Workers are more like to migrate a residence with a higher amenity value B_r, a higher benefit from local government goods $\frac{G_r}{R_r^\theta}$, and a lower per-unit price of floor space Q_r. In addition, there are two sources of bilateral determinants. The probability of choosing residence r decreases in the cost of migration D_{or}, but increases in the benefit of accessing the labor market discounted by commuting and job finding costs $\sum_{m=1}^{J} M_m \left(\frac{(1-\tau_m) w_m}{D_{om} D_{rm}} \right)^\epsilon$, which corresponds to the augmented labor market access $ALMA_{or}$. Using these conditional probabilities, migration market clearing condition requires that the number of workers who live in r is equal to the sum of workers migrating to r from all possible origins o:

$$R_r = \sum_{o=1}^{J} \pi_{r|o} R_o = \sum_{o=1}^{J} \frac{\left(\frac{\tilde{B}_r}{D_{or} Q_r^{1-\beta}} \left(\frac{G_r}{R_r^\theta} \right)^\lambda \right)^\epsilon ALMA_{or}}{\sum_{r'=1}^{J} \left(\frac{\tilde{B}_{r'}}{D_{or'} Q_{r'}^{1-\beta}} \left(\frac{G_{r'}}{R_{r'}^\theta} \right)^\lambda \right)^\epsilon ALMA_{or'}} R_o. \tag{1.19}$$

I derive the expression for the probability of commuting commuting to workplace m conditional on living in residence r. I take the ratio of the unconditional joint distribution of workers in terms of their residence and workplace to the the unconditional distribution of workers by residence as follows:

$$\pi_{m|r} = \frac{\sum_{o=1}^{J} \pi_{orm}}{\sum_{m'=1}^{J} \sum_{o'=1}^{J} \pi_{o'r'm'}} = \frac{\sum_{r0=1}^{J} \Phi_{orm}\pi_o/\Phi_o}{\sum_{m'=1}^{J} \sum_{o'=1}^{J} \Phi_{o'rm'}\pi_{o'}/\Phi_{o'}}$$
$$= \frac{\left(\frac{(1-\tau_m)\tilde{w}_m}{D_{rm}}\right)^{\epsilon} \sum_{o=1}^{J} \frac{\pi_o/\Phi_o}{(D_{or}D_{om})^{\epsilon}}}{\sum_{m'=1}^{J} \left(\frac{(1-\tau_{m'})\tilde{w}_{m'}}{D_{rm'}}\right)^{\epsilon} \underbrace{\sum_{o'=1}^{J} \frac{\pi_{o'}/\Phi_{o'}}{(D_{o'r}D_{o'm})^{\epsilon}}}_{AMMA_{rm'}}},$$

where the terms specific to current residence such as amenities, housing prices, and government goods are canceled out from the numerator and denominator. In line with intuition, workers are more likely to commute to places with higher returns $((1-\tau_m)\tilde{w}_m)^{\epsilon}$ net of commuting costs D_{rm}. Moreover, the conditional probability of commuting depends on how costly it is to migrate to residence r and find a job in workplace m, $\sum_{o=1}^{J} \frac{\pi_o/\Phi_o}{(D_{or}D_{om})^{\epsilon}}$, which corresponds to augmented migrant market access $AMMA$. Using these probabilities, I obtain the following expression:

$$L_m = \sum_{r=1}^{J} \pi_{m|r} R_r = \sum_{r=1}^{J} \frac{\left(\frac{(1-\tau_m)\tilde{w}_m}{D_{rm}}\right)^{\epsilon} AMMA_{rm}}{\sum_{m'=1}^{J} \left(\frac{(1-\tau_{m'})\tilde{w}_{m'}}{D_{rm'}}\right)^{\epsilon} AMMA_{rm'}} R_r, \tag{1.20}$$

where the number of workers employed in m is equated with the number of workers choosing to commute to m from all possible residences. I refer to this equation as the commuting market clearing condition.

Expected income of workers living in district r is equal to the sum of the after-tax wages in all possible workplace locations weighted by the conditional probabilities of commuting to those locations:

$$\mathbb{E}[(1-\tau_m)w_m|r] = \sum_{m=1}^{J} \frac{\left(\frac{(1-\tau_m)\tilde{w}_m}{D_{rm}}\right)^{\epsilon} AMMA_{rm}}{\sum_{m'=1}^{J} \left(\frac{(1-\tau_{m'})\tilde{w}_{m'}}{D_{rm'}}\right)^{\epsilon} AMMA_{rm'}} (1-\tau_m)w_m. \tag{1.21}$$

Expected income of workers are higher in places with lower costs of commuting D_{rm} as well as higher $AMMA_{rm}$, the indirect cost of commuting rising from the costs associated with migration and job finding. Because workers allocate $1-\beta$ fraction of their income to housing, the demand for residential floor space is given by

$$H_r^R = (1-\beta)\frac{\mathbb{E}[(1-\tau_m)w_m|r]R_r}{Q_r}. \tag{1.22}$$

Lastly, the population mobility implies that the ex-ante expected utility for each initial residence is the same across all possible residence-workplace pairs. That is,

$$\mathbb{E}[u_o] = \Gamma(\frac{\epsilon-1}{\epsilon})\Phi_o^{1/\epsilon} = \Gamma(\frac{\epsilon-1}{\epsilon})\left[\sum_{r'=1}^{J}\sum_{m'=1}^{J}\left(\frac{\tilde{B}_{r'}(1-\tau_{m'})\tilde{w}_{m'}}{D_{or'm'}Q_{r'}^{1-\beta}}\left(\frac{G_{r'}}{R_{r'}^{\theta}}\right)^{\lambda}\right)^{\epsilon}\right]^{1/\epsilon} \equiv \bar{u}_o, \quad (1.23)$$

where the expectation is taken over the distribution of the idiosyncratic component of utility.[48] I construct a measure of economy-wide welfare by taking the average of the expected utilities (1.23) weighted by the distribution of workers by their origins π_o: $\bar{u} = \sum_{o=1}^{J} \bar{u}_o \pi_o$. This measure corresponds to consumption equivalent worker welfare.

1.6.2 Production

The production of the tradable final good occurs under conditions of perfect competition and constant returns to scale. In particular, I assume that the production technology follows Cobb-Douglas as follows:

$$y_m = A_m L_m^{\alpha} {H_m^F}^{1-\alpha} \quad (1.24)$$

where A_m is final goods productivity; L_m is labor input; and H_m^F corresponds to a measure of floor space used commercially. Profit maximization under perfect competition implies that labor demand is high in places where productivity A_m is high; and wages w_m are lower in places with higher floor space available for commercial use H_m^F. This is captured in the labor demand as follows:

$$L_m = \left(\frac{\alpha A_m}{w_m}\right)^{\frac{1}{1-\alpha}} H_m^F. \quad (1.25)$$

The equilibrium wage equates the labor demand (1.25) to the labor supply (1.20) in each location. Similarly, the demand for floor space is given by

$$H_m^F = (\frac{(1-\alpha)A_m}{Q_m})^{\frac{1}{\alpha}} L_m. \quad (1.26)$$

The demand for floor space is high in a district with the low equilibrium floor space price Q_m, high productivity A_m, and measure of workers L_m.

[48] See Appendix 1.14.1 for the derivation of Equation (1.23).

1.6.3 Floor Space Market Clearing

There is a fixed floor space for each district H_j, which can be used residentially and commercially. Atomistic absentee landlords allocate ϑ_j fraction of $H_j$to commercial use and $1-\vartheta_j$ to residential use. Therefore, market clearing for residential floor space requires that the demand and supply of residential space are equal to each other (i.e., $H_j^F = (1-\vartheta_j)H_j$):

$$(1-\beta)\frac{\mathbb{E}[(1-\tau_m)w_m|r]R_j}{Q_j} = (1-\vartheta_j)H_j. \tag{1.27}$$

Commercial floor space market clearing requires that the demand for commercial floor space equals the supply of floor space allocated to commercial use (i.e., $H_j^R = \vartheta_j H_j$):

$$(\frac{(1-\alpha)A_j}{Q_j})^{\frac{1}{1-\alpha}}L_j = \vartheta_j H_j. \tag{1.28}$$

The setup of the floor space market in my model is consistent with the standard approach in the urban literature of assuming fixed supply (Rosen, 1979; Roback, 1982; Tsivanidis, 2019) and allowing residential and commercial uses (Ahlfeldt et al., 2015; Monte et al., 2018; Tsivanidis, 2019).[49]

1.6.4 National and Local Governments

Consistent with the national fiscal policies discussed in 1.2.3, I model how local government expenditures are determined. First, the national government determines a progressive income tax schedule $\tau(w)$, which is increasing in w, for all districts to levy their residents and collects the fraction of local tax revenue $1-\varsigma$ from each district. This means that ς fraction of total local tax revenue is kept locally, while $1-\varsigma$ fraction is delivered to the national government. I refer the parameter ς to as local-national revenue sharing. Also, without loss of generality, I express $\tau(w_m) = \tau_m$.

Second, the national government operates intergovernmental transfers to supplement tax revenues retained locally. It allocates χ fraction of the national tax revenue (or equiva-

[49] The choice to assume a fixed stock of floor space for each district is to focus on evaluating the consequence of spatial distribution of government spending. It is reasonable to assume that the total stock of floor space does not adjust instantly. While the total stock for each district is fixed, the model allows its allocation to residential and commercial uses to vary. As discussed in Redding and Rossi-Hansberg (2017), assuming absentee landlord following the urban economics literature does not allow the model to capture full general equilibrium effects. In addition, in my model, a single floor space price for each unit clears the floor space market clearing conditions for both the residential and commercial floor space markets. An extension to the model can be easily made to incorporate land use regulations limit the return to floor space allotted to commercial use as in Ahlfeldt et al. (2015) and Tsivanidis (2019).

lently, $(1-\varsigma)\chi$ fraction of total local tax revenue) for redistribution via intergovernmental transfers. Then, the national government determines the shares ς_j of the budget allotted for intergovernmental transfers to be delivered to each local governments such that $\varsigma_j \geq 0$ for all $j = 1, ..., J$ and $\sum_{j=1}^{S} \varsigma_j = 1$. I refer to $\{\varsigma_j\}_{j=1}^{J}$ as rules of redistribution. Lastly, the national government uses $(1-\varsigma)(1-\chi)$ fraction of total local tax revenue to provision national government goods and services such as national defense and diplomacy. I assume that national government goods and services benefit workers equally regardless of where workers live and work.

Given the national fiscal policies $\{\{\tau_m\}_{m=1}^{J}, \varsigma, \chi, \{\varsigma_j\}_{j=1}^{J}\}$, measure of workers living in j (R_j), conditional probabilities of commuting ($\{\pi_{m|j}\}_{m=1}^{J}$), and wages ($\{w_m\}_{m=1}^{J}$) determine local government budget in district j. A budget balancing equation of local government in district j is expressed as follows

$$G_j = \varsigma \underbrace{\sum_{m=1}^{J} \tau_m w_m \pi_{m|j} R_j}_{TR_j} + \varsigma_j (1-\varsigma)\chi \sum_{j'=1}^{J} TR_{j'}, \tag{1.29}$$

where $\sum_{m=1}^{J} \tau_m w_m \pi_{m|j} R_j$ is equal to local tax revenue collected from workers living in district j denoted by TR_j. Therefore, the first term corresponds to local tax revenue collected and retained by local government in district j. The second term is the amount of intergovernmental transfers from the national government, equal to the redistribution parameter for district j (ς_j) multiplied by the total budget allotted for intergovernmental transfers, $(1-\varsigma)\chi \sum_{j'=1}^{S} TR_{j'}$. The extent of fiscal decentralization is captured by $\tilde{\chi} = \varsigma + (1-\varsigma)\chi$, which corresponds to the fraction of total tax revenue spent locally.

Depending on the rules of redistribution, local government expenditure in a district may be greater if $\varsigma_j > TR_j / \sum_{j'=1}^{J} TR_{j'}$ or less than its contribution to intergovernmental transfers. In this sense, the spatial distribution of local government spending is considered as a consequence of transfers across districts. The redistribution mechanism described in this section has features that are structurally similar to a transfer scheme based on lump-sum tax and government spending laid out in Fajgelbaum and Gaubert (2018) and more broadly place-based policies (Glaeser and Gottlieb, 2008; Kline and Moretti, 2014).

1.6.5 General Equilibrium

Given vectors of exogenous location characteristics $\{T_j, M_j, B_j, A_j, d_{jk}, H_j\}$, initial distribution of workers $\{\pi_o\}$, total measure of workers R, national fiscal policies $\{\tau_j, \varsigma, \chi, \varsigma_j\}$, and model parameters $\{\alpha, \beta, \lambda, \theta, \kappa, \rho, \delta, \epsilon\}$, a general equilibrium of this economy is defined as a vector of endogenous objects $\{R_j, L_j, w_j, Q_j, \vartheta_j, G_j, \bar{u}_o\}$. These seven components of the equilibrium vector are determined by the migration market clearing (1.19), commuting market clearing (1.20), labor market clearing (1.25), floor space market clearing for residential and commercial uses (1.27 and 1.28), local government budget balancing equation (1.29), and population mobility (1.23).

1.7 Parameterization of the GE Model

So far, I have estimated one structural parameter governing the extent of rivalry associated with benefits from local government spending (θ) in Section 1.4.2 and five reduced-form elasticities: the elasticities of worker mobility to local government expenditure ($\lambda\epsilon$) and to home prices ($(1-\beta)\epsilon$) in Section 1.4.2 and the semi-elasticities of migration, commuting, and job finding with respect to distance ($\rho\epsilon$, $\kappa\epsilon$, $\delta\epsilon$) in Section 1.5.2. In this section, I discuss how I estimate the rest of the model parameters and recover unobserved local characteristics for year 2015.

1.7.1 Labor Share in Production and Housing Expenditure Share

First, the labor share (α=0.823) is estimated by computing average share of labor cost to the total costs across districts reported in Economic Census in 2015, consistent with the findings of Valentinyi and Herrendorf (2008). Second, I set housing expenditure $1-\beta$ equal to 0.15 to match the observed housing expenditure share based on Household Expenditure Survey in 2015. This value is corroborated with the reported value reported in OECD (2016).

1.7.2 National Fiscal Policy Parameters

The values of the national policy parameters are directly observed in a collection of laws governing local fiscal capacities (the Local Tax Act and the Local Subsidy Act). In 2015, $\varsigma = 9.1\%$ of local tax revenue retained after tax collection according to the Local Tax Act. The Local Subsidy Act allocates $\chi = 35\%$ of total local tax revenue delivered to the national government for redistribution. Because I observe the amount of intergovernmental transfers (IT_j) for each district, I recover the values for the redistribution parameters as follows:

$$\varsigma_j = \frac{IT_j}{\sum_{j'=1}^{J} IT_{j'}}. \tag{1.30}$$

The last national policy parameter of interest is the tax rates. The tax rates by income brackets are observed in the Income Tax Act as discussed in Section 1.2.3. However, the observed tax rates cannot be directly used because I do not observe the distribution of wages within each district, nor does the model feature wage dispersion within each locality. Without relying on the observed tax rates, I solve for tax rates τ_m by district based on the observed local tax revenue (LT_r), probability of commuting ($\pi_{m|r}$), wages (w_m), and number of workers by residence (R_r) by inverting the following system of equations:

$$\frac{1}{\varsigma}\begin{bmatrix} LT_{r=1} \\ \vdots \\ LT_{r=J} \end{bmatrix} = \left(\begin{bmatrix} \tau_{m=1} \\ \vdots \\ \tau_{m=J} \end{bmatrix} l_{1\times J} I_{J\times J} \begin{bmatrix} w_{m=1} \\ \vdots \\ w_{m=J} \end{bmatrix} \right)' \begin{bmatrix} \pi_{m=1|r=1} & \cdots & \pi_{m=J|r=1} \\ \vdots & \ddots & \vdots \\ \pi_{m=1|r=J} & \cdots & \pi_{m=J|r=J} \end{bmatrix} \begin{bmatrix} R_{r=1} \\ \vdots \\ R_{r=J} \end{bmatrix}, \tag{1.31}$$

where $l_{1\times J}$ is a vector with all of its elements equal to 1. Finally, I simplify the tax rates to a single index ($\tau_m = \tau\ \forall m$) and calibrate the simplified tax rate equal to 0.245, the average tax rates from the inversion weighted by number of workers.[50]

1.7.3 Recovery of Unobserved Local Characteristics

1.7.3.1 Local Productivity

I recover the values for local productivity using the observed wages and floor space prices. To satisfy the profit maximization and zero profit conditions, equilibrium floor space prices must satisfy:

$$Q_j = (1-\alpha)\left(\frac{\alpha}{w_j}\right)^{\frac{\alpha}{1-\alpha}} A_j^{\frac{1}{1-\alpha}}. \tag{1.32}$$

Therefore, given the observed data on wages and floor space prices in 2015 and the parameter value of α, I can recover A_j for each district using the equilibrium condition above (1.32). Figure 1.12 in Appendix 1.13.2 plots the spatial distribution of the recovered values of local productivity. The greater Seoul area, the Northwestern part of South Korea, has relatively greater values of productivity, as well as the some of the coastal districts with ports (e.g., the greater Busan area covering the Southeastern coast) consistent with coastal and port advantages studied in Balboni (2019) and Ducruet et al. (2019).

[50] The counterfactual policy experiments concern with changes in the spatial distribution of spending due to changes in the intensity of fiscal decentralization and redistribution whiling holding the nationally determined tax rates fixed.

1.7.3.2 Fréchet Shape Parameter

I estimate the Fréchet shape parameter, which is equivalent to the elasticity of worker mobility with respect to wage. I begin by deriving the expression for the probabilities of working in m conditional on living in r and having moved from o:

$$\pi_{m|ro} = \frac{\pi_{orm}}{\sum_{m'=1}^{J} \pi_{orm'}} = \frac{\Phi_{orm}}{\sum_{m'=1}^{J} \Phi_{orm'}} = \frac{\frac{\tilde{w}_m^{\epsilon}}{\exp(\kappa\epsilon d_{rm} + \delta\epsilon d_{om})}}{\sum_{m'=1}^{J} \frac{\tilde{w}_{m'}^{\epsilon}}{\exp(\kappa\epsilon d_{rm'} + \delta\epsilon d_{om'})}}. \tag{1.33}$$

I define a composite referred to as adjusted wages $\omega_j = \tilde{w}_j^{\epsilon} = M_j w_m^{\epsilon}$. I rewrite the above equation using adjusted wages and take the log transformation of both sides. Using my estimates of $\kappa\epsilon = 0.045$ and $\delta\epsilon = 0.016$ and rearranging such that left hand side consists of only observables, I obtain the following expression:

$$\ln \pi_{m|ro} + \kappa\epsilon d_{rm} + \delta\epsilon d_{om} = -\ln \sum_{m'=1}^{J} \frac{\omega_{m'}}{\exp(\kappa\epsilon d_{rm'} + \delta\epsilon d_{om'})} + \ln \omega_m, \tag{1.34}$$

where I treat $\kappa\epsilon d_{rm} + \delta\epsilon d_{om}$ as data and I observe $\ln \pi_{m|ro}$. The left hand side altogether can be decomposed into two parts: the first term that varies at the current residence and origin level and the second term that varies at the workplace level. Introducing stochastic errors to Equation (1.34), I regress the left hand side on the pairwise fixed effects of current residence and origin and the workplace fixed effects. Then, I recover the values of log adjusted wages from the estimated workplace fixed effects. Note that these values are determined independent of ϵ based on the observed distribution of workers and the costs of commuting and job finding.[51]

The parameter ϵ controls the variance of log adjusted wages ($\ln \omega_j$) relative to the variance of log observed wages ($\ln w_m^{\epsilon}$). That is, $\sigma^2_{\ln w_j} = \frac{1}{\epsilon^2}\sigma^2_{\ln \omega_j}$ because the parameters M_j are deterministic. Therefore, I estimate the value of ϵ by taking the ratio of the standard deviations of log adjusted wages and log wages in the data after normalizing both to have geometric mean equal to 1. The resulting value of ϵ is equal to 3.54; this means that the worker mobility increases by 3.54 percent for a 1 percent increase in wages.[52]

[51] In Appendix 1.13.2, I examine the relationship between the recovered values of local productivity and number of firms in Figure 1.13. Panel (a) shows that districts with higher productivity values have higher number of firms. In Panel (b), productivity is positively correlated with number of firms which discharge wastewater.

[52] For inference, I randomly sample 111 observations from log observed wages and log adjusted wages and compute ϵ. I repeat this process a number of times, e.g., 1 billion times, to obtain the distribution of the estimator for ϵ. Given the estimated standard error equal to 0.102. I reject the null hypothesis that $\epsilon = 0$ at the 99% confidence level. The result is robust to different sample sizes (25, 50, 100, 150 and 200).

There are several other papers which estimate the same parameter. Defining spatial units as U.S. counties from 2006 to 2010, Monte et al. (2018) finds a point estimate of the shape parameter equal to 3.3, while Ahlfeldt et al. (2015) estimate its value equal to 6.83 based on the inter-district commuting patterns in the city of Berlin in 2008. My estimate falls within the responsible range of the existing estimates in the literature. With the estimated value of ϵ, I recover the structural parameters $(\lambda, \rho, \kappa, \rho)$ from the estimated reduced-form elasticities ($\lambda = \tilde{\lambda}/\epsilon = 0.30$, $\rho = \tilde{\rho}/\epsilon = 0.009$, $\kappa = \tilde{\kappa}/\epsilon = 0.013$, $\delta = \tilde{\delta}/\epsilon = 0.005$). Furthermore, the estimated elasticity of worker mobility with respect to floor space prices reported in Table 1.3 is equal to $(1-\beta)\epsilon$. Based on the estimate of $\epsilon = 3.54$, the implied value of $1-\beta$ is equal to 0.14, close to the expenditure share estimated using the Household Expenditure Survey in Section 1.7.1.

Based on the structural value of how much people of local government spending ($\lambda = 0.3$), I obtain the valuation of local government spending by computing the compensating variation. At the median values of per-capita local government spending (7,302 USD) and household income (18,180 USD) in 2015, workers are willing to give up 75 cent for a dollar increase in per-capita local government expenditure in their residence.

1.7.3.3 Adjusted Local Amenities

I recover adjusted amenity for each residence that rationalizes the observed spatial distribution of workers. Similarly to the process described when recovering the adjusted wages, I begin by deriving the expression for the conditional distribution of workers by their residences on workplace location and previous residence based on the gravity equation (1.5):

$$\pi_{r|mo} = \frac{\pi_{orm}}{\sum_{r'=1}^{J} \pi_{or'm}} = \frac{\Phi_{orm}}{\sum_{r'=1}^{J} \Phi_{or'm}} = \frac{\frac{\tilde{B}_r^{\epsilon} G_r^{\lambda\epsilon}}{\exp(\kappa\epsilon d_{rm} + \rho\epsilon d_{or}) Q_r^{(1-\beta)\epsilon} R_r^{\theta\lambda\epsilon}}}{\sum_{r'=1}^{J} \frac{\tilde{B}_{r'}^{\epsilon} G_{r'}^{\lambda\epsilon}}{\exp(\kappa\epsilon d_{r'm} + \rho\epsilon d_{or'}) Q_{r'}^{(1-\beta)\epsilon} R_{r'}^{\theta\lambda\epsilon}}}. \tag{1.35}$$

I take the log transformation of both sides of Equation (1.35) and rearrange such that left hand side only consists of observables:

$$\ln \pi_{r|mo} - \ln \frac{G_r^{\lambda\epsilon}}{\exp(\kappa\epsilon d_{rm} + \rho\epsilon d_{or}) Q_r^{(1-\beta)\epsilon} R_r^{\theta\lambda\epsilon}} = -\ln \sum_{r'=1}^{J} \frac{\tilde{B}_{r'}^{\epsilon} G_{r'}^{\lambda\epsilon}}{\exp(\kappa\epsilon d_{r'm} + \rho\epsilon d_{or'}) Q_{r'}^{(1-\beta)\epsilon} R_{r'}^{\theta\lambda\epsilon}} + \ln \tilde{B}_r^{\epsilon}, \tag{1.36}$$

where I treat the second term in the left hand side as data given the parameter values. Introducing stochastic errors, I regress the left hand side on the pairwise fixed effects of workplace and origin and the residence fixed effects. Then, I recover the values of log adjusted amenities from the estimated residence fixed effects (up to scale). Figure 1.14 in Appendix 1.13.3 plots the spatial distribution of adjusted amenities. The metropolitan areas tend to have relatively higher amenity values, reflecting urban amenities. Also, the amenities are higher in the coastal areas, especially the coastal districts in the East and South.[53]

1.7.4 Non-targeted Moments

I evaluate how well the model predicts the non-targeted moments. First, I compare the observed data on number of workers by employment location to the model prediction in Panel (a) and (b) of Figure 1.5. The two variables have a coefficient correlation of 0.94 with a slope equal to 0.91 in Panel (a). The estimated slope in Panel (a) as well as the comparison of the cumulative distribution functions in Panel (b) suggest that the model performs well in explaining the spatial distribution of workers.

Second, in Panel (c) and (d), I compare the observed local tax revenue to the model-implied local tax revenue by residence. There is a strong positive correlation between the data and the model-implied local tax revenues with a value of 0.92 and an estimated slope of 0.95. In addition, I plot the cumulative distribution functions of the data on local tax revenue and the model-counterpart. Local government spending is equal to the sum of a fixed fraction of local tax revenues and the intergovernmental transfers, last of which my calibration matches. Therefore, Panel (c) and (d) show that the model explains the spatial distribution of local government spending well.

Third, I verify the model prediction on residential floor space. Panel (e) and (f) compare the residential floor spaces predicted by the model to the observed area of land used for residential purposes measured in $1000m^2$ from the Land Use Statistics in 2015. The correlation coefficient of the two variables is 0.52 and the estimated slope is equal to 0.97. While strong, the relationship between the data and the model-implied values has a relatively low correlation coefficient. This is because the observed data measures total land area used residentially, which does not take the ratio of floor space to land area into

[53] In Appendix 1.13.3, I assess the relationship between the recovered amenities and local outcomes which proxy quality of life at the residential locations (Desmet and Rossi-Hansberg, 2013). Panel (a) of Figure 1.15 shows that amenities are higher in places with fewer number of firms discharging wastewater. In Panel (b), residences with lower suicide rates tend to have higher amenities. Lastly, there is a negative correlation between divorce rates and amenities in Panel (c). While I do not formally investigate the relationship between weather and the recovered amenity values as in (Rappaport, 2007), I can infer that nice weather is positively correlated with the recovered amenities because coastal areas tend to have mild weather in Summer and Winter relative to inland districts. Therefore, proximity to the ocean in the coastal districts and its positive relationship with nice weather make coastal districts relatively more attractive.

account. Despite the sources of measurement error, the model performs well in capturing residential floor spaces.

1.8 Counterfactual Policy Experiments

In this section, I quantify the welfare consequences of the spatial distribution of local government spending. In particular, I vary the extent of redistribution while holding the rules of redistribution and the extent of fiscal decentralization constant.

1.8.1 Determinants of Rules of Redistribution

The primary objective of the Local Subsidy Act, which determines the rules of redistribution, is to promote equitable economic growth across localities. As a result, the rules of redistribution is expected to favor residences which are intrinsically less attractive to live (low values of $\tilde{B}_j$) and fiscally weak (low TR_j) to promote economic growth in these districts. It is important to understand the determinants of rules of redistribution because I conduct counterfactual policy experiments while holding the observed rules of redistribution fixed in the subsequent section. I formally study the determinants of the rules of redistribution observed in 2015 in a regression framework. To do so, I regress the log of the observed rules of redistribution $\ln \varsigma_j$ on the log of residential density R_j, recovered amenity values $\tilde{B}_r$, local productivity A_j, and employment density L_j. Table 1.8 summarizes the estimation results.

In Column (1), the coefficient in front of the log residential density is positive and statistically significant. The result implies that the rules of redistribution is higher in places with higher population density conditional on the geographical area. Introducing the log recovered values of adjusted amenities and productivity in Column (2) and then in Column (3), I find that districts with higher amenity values and productivity receives smaller share of intergovernmental transfers. Lastly, in Column (4), I find that the employment density of a residence does not affect the rules of redistribution.

1.8.2 Welfare Consequences of Redistribution

In this section, I conduct a series of counterfactual policy experiments in which I vary the extent of redistribution. Throughout the exercises, I hold the extent of fiscal decentralization (i.e., the fraction of total tax revenue spent locally) constant at the level observed in 2015 $\tilde{\chi} = 0.4$ as well as the rules of redistribution $\{\varsigma_j\}_{j=1}^{S}$. In each of counterfactuals, I consider varying extent of redistribution denoted by $\tilde{\varsigma}$, which varies from 0 up to $\tilde{\chi}$. Local government spending is expressed as follows:

$$G_j = (\tilde{\chi} - \tilde{\varsigma})TR_j + \varsigma_j \tilde{\varsigma} \sum_{j'=1}^{S} TR_{j'} \tag{1.37}$$

If $\tilde{\varsigma} = 0$, local government spending solely depends on local tax revenue. In the other extreme in which $\tilde{\varsigma} = \tilde{\chi}$, intergovernmental transfers completely determine local government expenditures. The observed extent of redistribution is 0.3, which I consider a baseline.

Figure 1.6 plots the changes in the aggregate welfare of workers $\bar{u}$ as defined in Section 1.6.1 relative to the baseline level ($\tilde{\varsigma} = 30\%$). When the redistributive intergovernmental transfers are completely eliminated and local government spending is determined solely based on local tax revenue ($\tilde{\varsigma} = 0\%$), the aggregate welfare of workers decrease by 1.2 percent. In the other extreme case in which local government spending is completely determined by intergovernmental transfers ($\tilde{\varsigma} = 40\%$), the aggregate welfare also decreases by 0.3 percent. Considering the varying extent of redistribution (with an increment of 5 percentage points), I find that the aggregate welfare is maximized when the extent of redistribution is equal to 20 percent. This implies that by lowering the extent of redistribution observed in 2015 by 10 percentage points, the aggregate welfare of workers would reach its highest, which is 0.12 percent higher than the baseline level.

The extent of redistribution controls the trade-offs between two types of fiscal spillovers. In districts that are net contributors to redistribution, a dollar tax contribution of a resident is shared with all the other residents living in the same district, but also with other workers living in districts that are net receivers.[54] Therefore, in the presence of redistributive intergovernmental transfers, there are two sources of fiscal spillovers: intra-district and inter-district. The size of intra-district fiscal spillover decreases in the extent of redistribution. It is also necessarily the case that the size of inter-district fiscal spillover becomes larger as the extent of redistribution increases.

Therefore, the welfare changes summarized in Figure 1.6 are the consequences of changes in the extents of intra- and inter-district fiscal spillovers. On the one hand, when the extent of redistribution is greater than 20 percent, inter-district spillover serves as a primary source of inefficiency. In this case, intergovernmental transfers raise local

[54] To help understand the types of spillovers, it is important to reiterate two important characteristics of local government spending in South Korea and more broadly local public finance. First, local government goods and services are not fully rival (i.e., $\theta < 1$). Second, due to redistributive intergovernmental transfers, how much is transferred from districts that are fiscally strong (net contributors) to those with weak fiscal capacities (net receivers) increases in the extent of redistribution. Higher the extent of redistribution, larger the fraction of my tax contribution diverted for redistribution.

government expenditures in net-receiving districts by drawing expenditures from net-contributing districts. In response, workers are attracted to and move to these places which have become less undesirable. On the other hand, when the extent of redistribution is less than 20 percent, intra-district spillover is responsible for lowering the overall welfare. Similarly, in this case, districts that are fiscally strong would attract additional residents from the tax contributions of fellow residents shared within each district.

Figure 1.7 shows that the extent of fiscal spillovers is minimized when the extent of redistribution is equal to 20 percent. I construct a measure for the extent of fiscal spillovers $\tilde{\varsigma}$ by computing the standard deviation of local government goods and services net of worker tax contribution (i.e., how much extra benefit workers enjoy due to spillovers) for each counterfactual. This measure gauges the dispersion of external benefits of local government spending from intra- and inter-district spillovers. Higher the dispersion, higher the incentives for the workers to reallocate. At the optimum level of redistribution at 20 percent, the extent of fiscal spillovers is reduced by 20 percent.

Lastly, I conduct the same set of counterfactual policy experiments based on two different restrictions commonly imposed in the literature on migration and commuting. First, the migration literature assumes that workers live and work in the same location. I set the semi-elasticity of commuting with respect to commuting distance $\kappa\epsilon$ equal to infinity, the semi-elasticity of migration with respect to migration distance $\rho\epsilon$ equal to 0.007, and the semi-elasticity of job finding to its distance equal to 0. Second, the commuting literature assumes costless migration. Likewise, I assume that the distance-elasticities of migration and job search equal to zero and the semi-elasticity of commuting to commuting distance equal to 0.074 and compute the counterfactual outcomes. Then, for each of two sets of redistribution separately, I solve for the new equilibrium and compute counterfactual changes in the aggregate worker welfare under varying extents of redistribution.

In Panel (a) of Figure 1.8, I plot the welfare changes relative to the baseline in 2015 assuming no inter-district commuting. If workers are not allowed to commute outside of districts, eliminating redistribution altogether leads in a higher welfare loss of about 2 percent. Furthermore, the optimal extent of redistribution is higher at a level close to 30 percent. Workers are not able to access districts with higher productivity without moving into these districts. Then, workers agglomerate in these districts, contributing to increasing intra-district fiscal spillover. As a result, there is a demand for greater redistribution.

Panel (b) of Figure 1.8 plots the changes in the worker welfare under the assumption of no migration and job finding costs. While not optimal, eliminating redistributive intergovernmental transfers lead to a sizable increase in welfare by about 2.3 percent. This implies that the need for redistribution across districts is small when workers can migrate across districts freely. With no migration and job finding costs, it becomes easier for workers to access districts with higher productivity. At the same time, in the presence of redistributive intergovernmental transfers, workers find it profitable to reside in net-receiving districts with positive inter-district fiscal spillovers at the expense of longer commute because they benefit from local government goods and services more than their tax contribution. Therefore, with no migration and job finding costs, lowering the extent of redistribution increases the overall efficiency of the economy.

1.9 Conclusion

In this paper, I make three contributions to our understanding of local provision of government goods and services as a determinant of the spatial distribution of workers. First, I present a quantitative general equilibrium in which workers make both migration and commuting decisions, which have been traditionally studied separately. The key prediction of the model is a gravity equation summarizing the distribution of workers in terms of three locations: previous residence, current residence, and workplace.

Second, I combine the framework with the quasi-natural experiment leading to plausibly exogenous variation in local spending and estimate the key reduced-form elasticities of worker mobility with respect to local government expenditure, residential density, and home prices. In addition, I estimate the elasticities of worker mobility with respect to spatial frictions (migration, commuting, and job finding). The key finding is that the marginal valuation of local government spending is equal to 75 cents of after-tax income. I show that there are large biases when the elasticity of migration with respect to distance is estimated without accounting for commuting patterns and vice versa. The results altogether show that where workers lived before matters for not only where they live today, but only where they presently work.

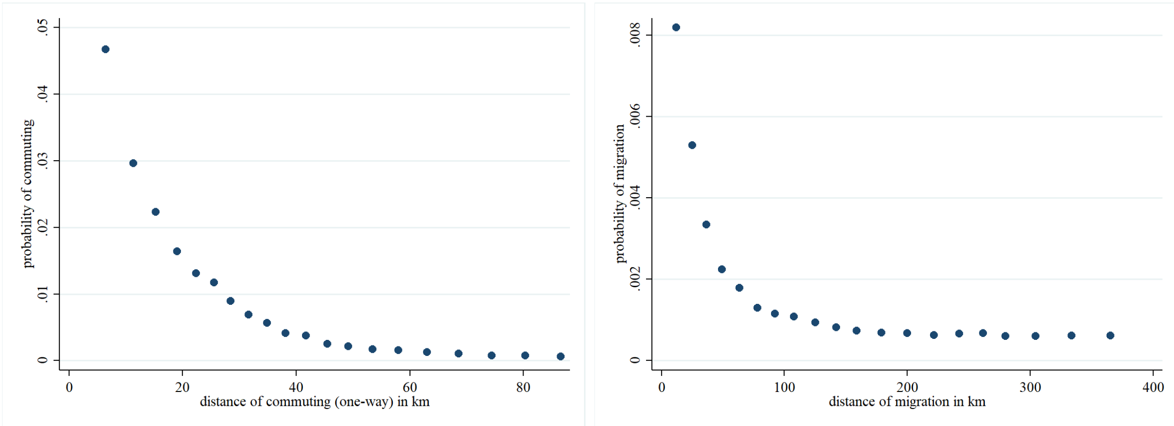
Third, based on counterfactual policy experiments, I show that there exists a fiscal arrangement (local taxation vs. intergovernmental transfers), which would maximize the overall welfare of workers. I discuss how the optimal mix of local taxation and intergovernmental transfers balances the trade-offs between the extents of intra-regional and inter-regional fiscal spillovers. The results suggest that reducing the current extent of redistribution observed in South Korea, thereby allowing local governments to rely more on their local income tax, would increase the overall efficiency. Furthermore, spatial frictions and their effects on worker mobility are important in determining an optimal fiscal arrangement. If workers are assumed to live and work in the same location (the key assumption in the migration literature), the importance of redistributive intergovernmental transfers are overemphasized. However, if workers can migrate without any costs (the key assumption in the commuting literature), the importance of local taxation (less redistribution) is overemphasized.

Overall, I find that it is crucial to account for both margins of mobility (i.e., migration and commuting) not only to understand the determinants and their effects on the spatial distribution of workers and more broadly economic activity, but also to inform policy makers of the welfare consequences from the spatial distribution of local government spending.

1.10 Figures and Tables

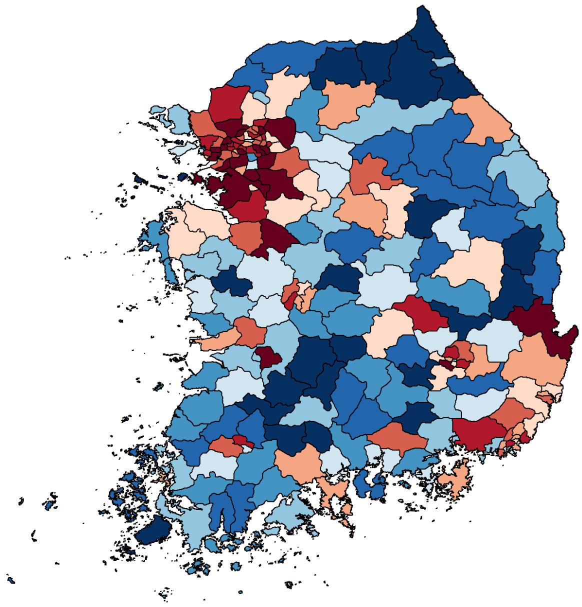
Figure 1.1: Commuting and Migration Patterns vs. Distance

(a) Commuting **(b)** Migration

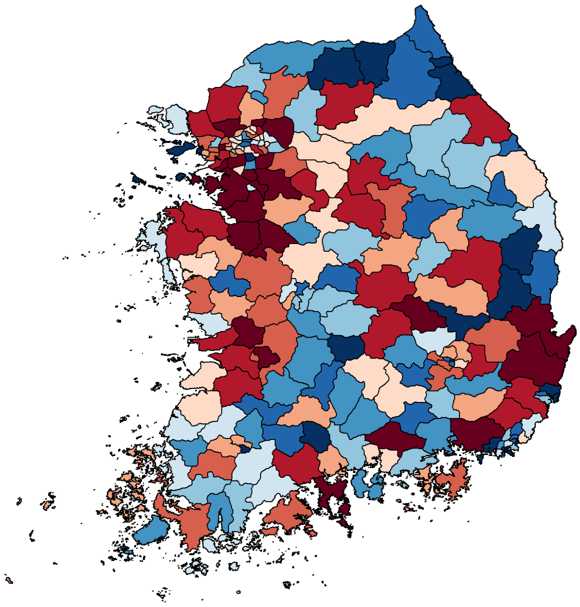


Notes: This figure shows that the probabilities of commuting shown in Panel (a) and the probabilities of migration shown in Panel (b) decrease as distances of commuting and migration increase. The probabilities of commuting and migration are computed using the Population Census of South Korea (2005, 2010, and 2015). Each point corresponds to 5 percentiles of commuting and migration distances.

Figure 1.2: Spatial Distribution of Residential Density and Local Government Spending



(a) Residential Density

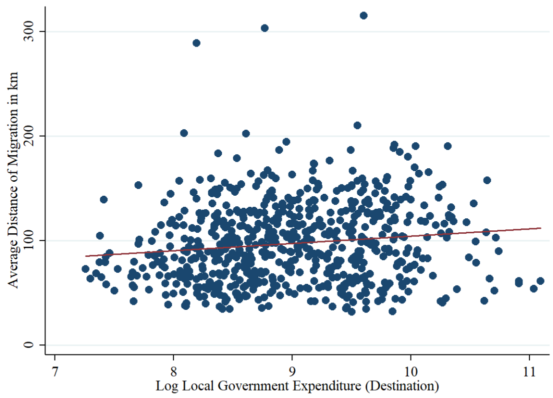


(b) Local Government Spending

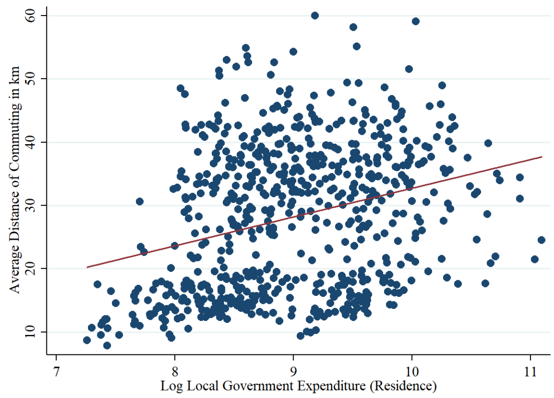
Notes: The figure on the left plots the spatial distribution of workers in terms of their residences in 2015. The figure on the right plots the spatial distribution of local government spending in 2015. Red (blue) districts indicate higher (lower) values.

Figure 1.3: Workers appear willing to migrate/commute longer with higher government spending and lower housing prices

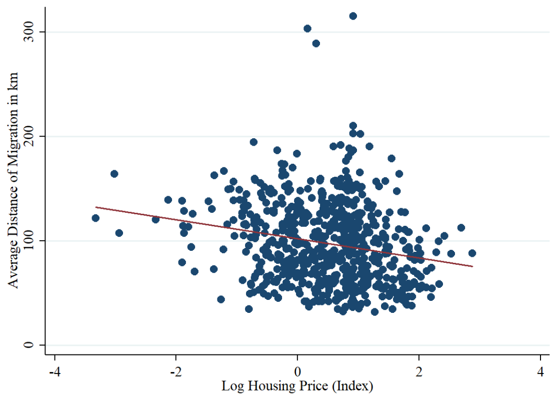
(a) Migration Distance vs. Local Gov't Spending



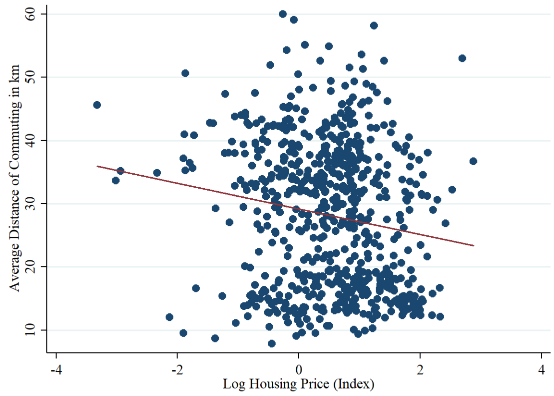
(b) Commute Distance vs. Local Gov't Spending



(c) Migration Distance vs. Home Prices

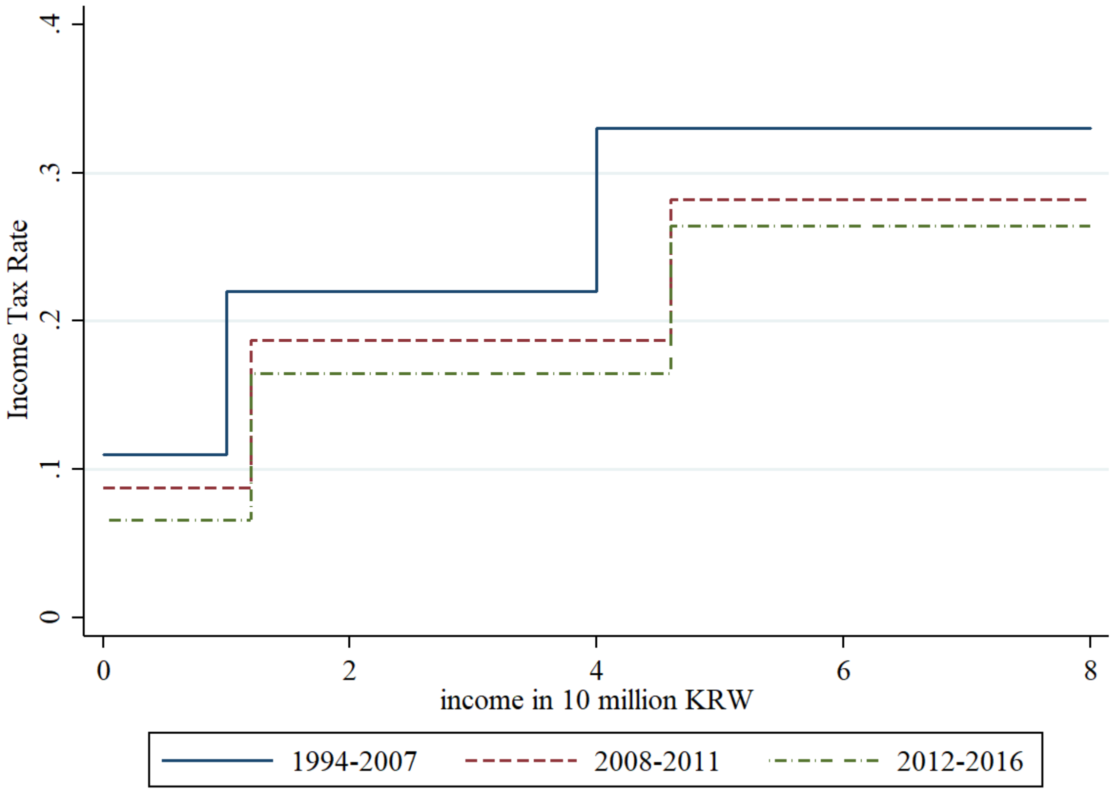


(d) Commute Distance vs. Housing Prices



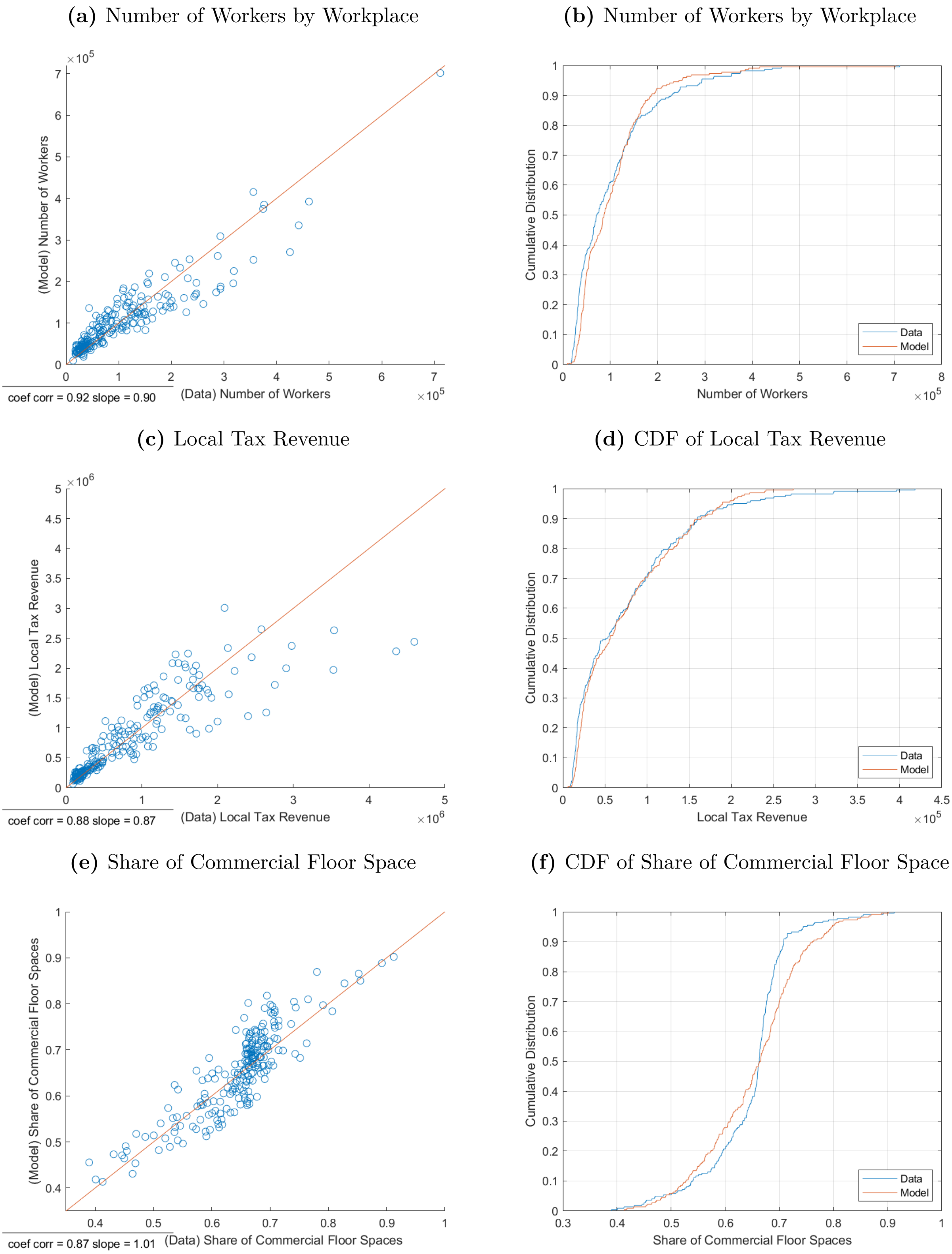
Notes: This figure shows the raw correlation between how far workers migrate and commute and local government spending/Housing prices. Each observation is a district-year pair. The figures in the left plot the average distance that residents have migrated over the past 5 years against local government spending in Panel (a) and against home prices in Panel (c). The figure in the right plot the average distance of commuting for a resident for each district against local government spending in Panel (b) and against the home prices in Panel (d).

Figure 1.4: Marginal Income Tax Rates before and after 2008 and 2012



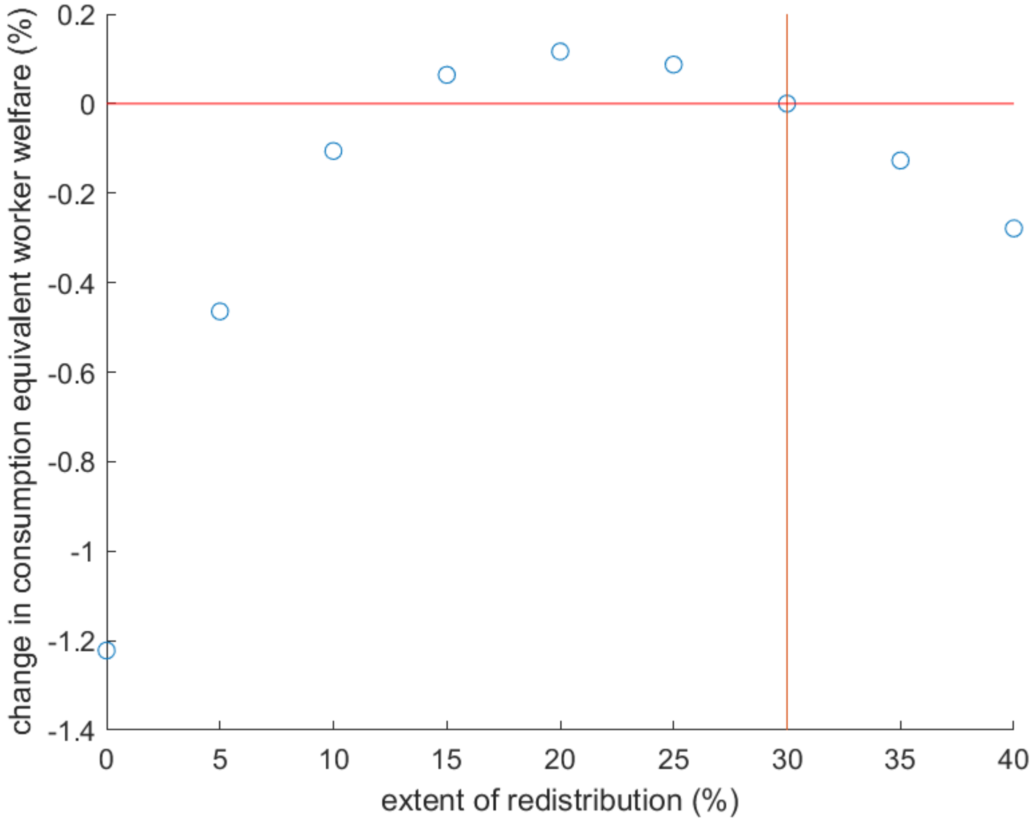
Notes: This figure plots the progressive income tax rates against income measured in 10 million KRW (approximately 10,000 USD) before and after the two episodes of national tax policy reforms in 2008 and 2012. The national income tax rates are outlined in the Income Tax Act. Note the median after-tax income in South Korea in 2015 is 18,180 USD.

Figure 1.5: Over-identifying Moments: Model vs. Data



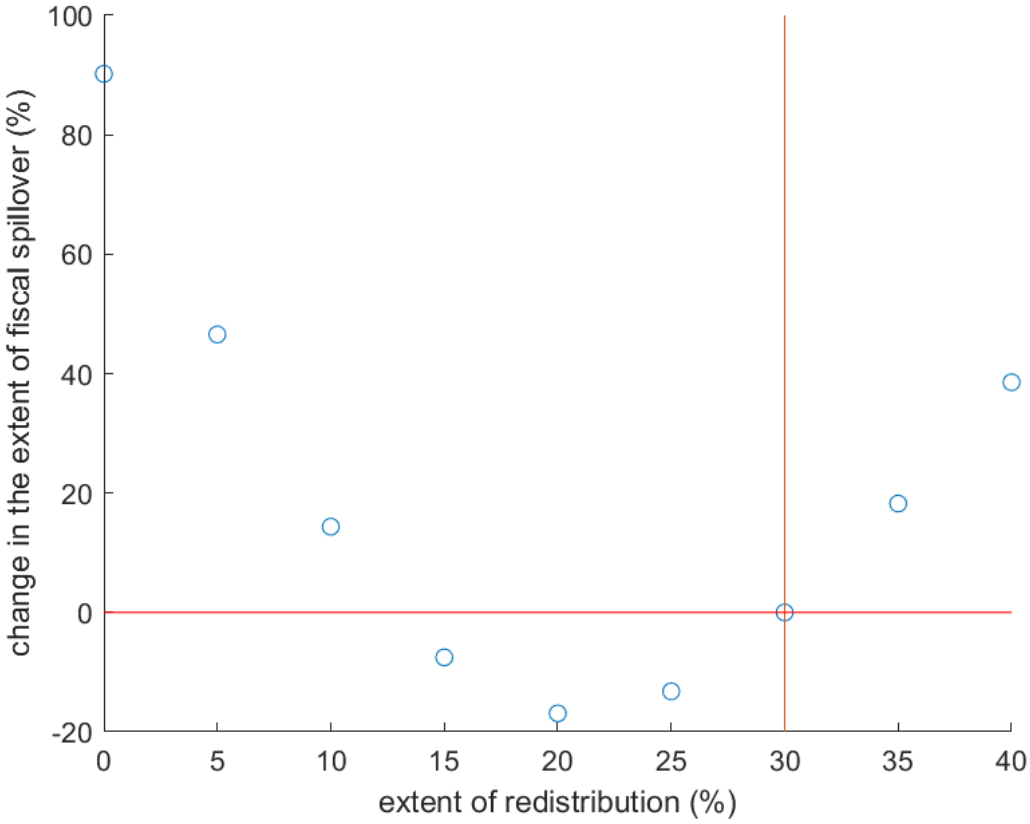
Notes: This figure compares 2015 data with model predictions of non-targeted moments. Panel (a) and (b) plot the spatial distribution of workers by employment location. Panel (c) and (d) plot local tax revenues collected at each residence measured in 1 million KRW. Panel (e) and (f) plot the shares of commercial floor space. The straight lines in Panel (a), (c), and (e) are 45 degree lines.

Figure 1.6: Aggregate Welfare Changes and Redistribution



Notes: This figure plots the changes in the aggregate consumption equivalent worker welfare relative to the welfare level in baseline in which the extent of redistribution is equal to 0.3, the observed level in 2015.

Figure 1.7: Changes Extent of Fiscal Spillover Changes and Redistribution

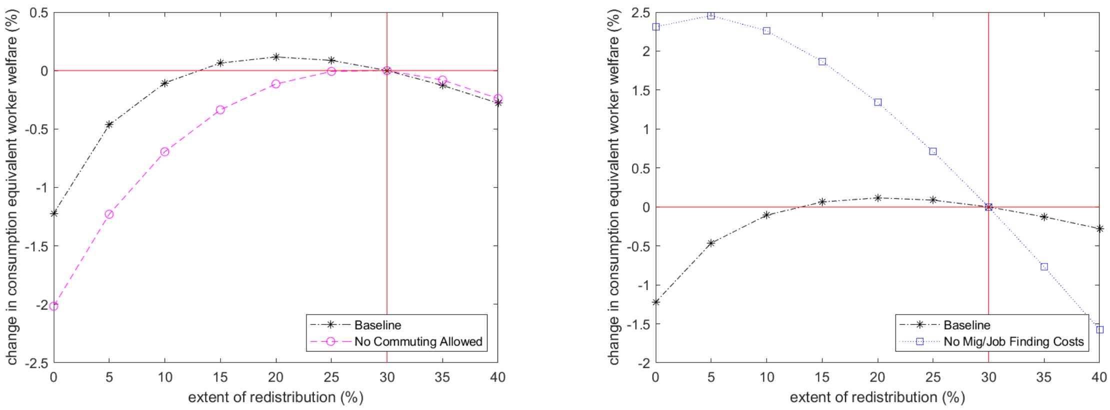


Notes: This figure plots the changes in extent of fiscal spillover to the baseline level in which the extent of redistribution is equal to 0.3, the observed level in 2015. The extent of fiscal spillover measures the dispersion of local government goods and services net of individual tax contribution (i.e., how much extra benefit workers enjoy due to spillovers) across districts.

Figure 1.8: Aggregate Welfare Changes under Alternative Assumptions

(a) Prohibitively Costly Commuting

(b) Costless Migration and Job Finding



Notes: In this figure, I plot the changes in the consumption equivalent welfare of workers based on two alternative assumptions about spatial frictions. First, I follow the common spatial redistribution imposed in the migration literature (i.e., workers cannot work outside of their district of residence). I solve for a new equilibrium for 2015 assuming the distance elasticity of migration equal to 0.007 as in Column (5) of Table 1.4, of commuting equal to ∞, and of job search equal to 0. I compute the counterfactual outcomes and plot the changes in worker welfare relative to 2015 (extent of redistribution = 30%) in Panel (a). Second, I follow the common spatial redistribution imposed in the commuting literature (i.e., there is no bilateral cost of commuting and job search). I solve for a new equilibrium for 2015 assuming the distance elasticity of migration equal to 0, of commuting equal to 0.074 as in Column (5) of Table 1.5, and of job search equal to 0. I compute the counterfactual outcomes and plot the changes in worker welfare relative to 2015 (extent of redistribution = 30 percent) in Panel (b).

Table 1.1: Summary Statistics

Variable	(1) Observation	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max
		A. Commuting Patterns			
Commuters from Residence	666	0.274	0.236	0	0.773
Commuters to Workplace	666	0.298	0.210	0	0.916
		B. Migration Patterns			
Migrants to Residence	666	0.187	0.079	0.053	0.559
Out-Migrants from Residence	666	0.180	0.071	0.048	0.443
		C. Local Government Budget			
Total Local Expenditure	666	362,785	233,524	59,614	1,881,082
Per-Capita Local Expenditure	666	7.638	5.752	0.904	29.622
Local Income Tax Revenue	666	64,067	95,793	4,020	779,143
Intergovernmental Transfers	666	242,517	128,348	109,239	799,009

Notes: In this table, I report summary statistics computed based on 222 districts in 2005, 2010, and 2015. The data used for Panel A and B is constructed from the Population Census of South Korea. Variable *Commuters from Residence* measures the fraction of residents commuting outside of their district of residency. Variable *Commuters to Workplace* measures the fraction of workers employed in a district who commute from other districts. Similarly, Variable *Migrants to Residence* and *Out-Migrants from Residence* measure the fraction of residents in a district who moved in within 5 years and the fraction of residents who moved out of a district within 5 years. Panel C is computed using the Yearbook of Local Public Finance data. The unit for the values reported in Panel (c) is 1 million KRW (approximately 1,000 USD). See Section 1.2.1 for the details on the data sources.

Table 1.2: (OLS) Elasticities of Worker Mobility with respect to Local Government Goods

	(1)	(2)	(3)	(4)
Dependent Variable:	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$
Local Government Expenditure, $\ln G_{r,t}$ ($\beta_G = \lambda\epsilon$)	-0.231***	0.0965**	0.101	0.0957***
	(0.0149)	(0.0405)	(0.141)	(0.0299)
Number of Households, $\ln R_{r,t}$ ($\beta_R = \theta\lambda\epsilon$)	0.120***	0.0608**	0.297	0.590***
	(0.0125)	(0.0268)	(0.218)	(0.0522)
Floor Space Price, $\ln Q_{r,t}$ ($\beta_Q = (1-\beta)\epsilon$)	-0.0416***	-0.0101	0.00802	-0.00148
	(0.0129)	(0.0334)	(0.0336)	(0.00653)
Observations	258,323	258,323	258,323	258,323
Fixed Effects:				
Job Finding Pair ×Year ($\phi_{om,t}$)	N	Y	Y	Y
Migration Pair (ϕ_{or})	N	N	Y	Y
Commuting Pair (ϕ_{rm})	N	N	N	Y

Notes: In this table, I report the OLS estimates of elasticities of worker's mobility to local government expenditure and resident population levels based on Equation 1.6, starting with a simple estimate without any fixed effects in Column (1) and gradually adding the fixed effects discussed in Section 1.4.1.1. Column (4) corresponds to Equation 1.6 with the full set of fixed effects. The sample is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location. Robust standard errors in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, and a triplet of previous and current residences and workplace: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

Table 1.3: (2SLS) Elasticities of Worker Mobility with respect to Local Government Goods

	(1)	(2)	(3)	(4)	(5)
	OLS	First Stage	First Stage	First Stage	2SLS
Dependent Variable:	$\ln \pi_{orm,t}$	$\ln G_{r,t}$	$\ln R_{r,t}$	$\ln Q_{r,t}$	$\ln \pi_{orm,t}$
Local Government Expenditure, $\ln G_{r,t}$ ($\beta_G = \lambda\epsilon$)	0.0957***				1.072***
	(0.0299)				(0.387)
Number of Households, $\ln R_{r,t}$ ($-\beta_R = -\theta\lambda\epsilon$)	0.590***				-0.844
	(0.0522)				(0.622)
Floor Space Prices, $\ln Q_{r,t}$ ($-\beta_Q = -(1-\beta)\epsilon$)	-0.00148				-0.490***
	(0.00653)				(0.067)
Predicted Tax Contribution (low), $IV_{r,t}^{low}$		13.26***	5.592***	60.23***	
		(1.518)	(0.246)	(6.851)	
Predicted Tax Contribution (high), $IV_{r,t}^{high}$		13.921***	6.761***	21.30***	
		(0.833)	(0.246)	(3.257)	
Number of Households 30 years ago, $IV_{r,t}^{R}$		0.028***	-0.014**	0.110***	
		(0.009)	(0.007)	(0.030)	
SW F-stat		19.01	16.87	25.66	
$\hat{\theta}$	-6.164**				0.787**
	(2.176)				(0.315)

Notes: In this table, I compare the OLS estimates and 2SLS estimates of elasticities of worker's mobility to local government expenditure and resident population levels based on Equation 1.6. The number of observations is 258,323, the same across columns. Column (1) is identical to Column (4) in Table 1.2. Column (2) and Column (3) report the first stage results. The 2SLS estimates are reported in Column (4). Across columns, the full set of fixed effects as discussed in Section 1.4.1.1 are included. The sample ($N = 258,323$) is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location. Robust standard errors for Column (1), (2), and (3) and bootstrapped (20,000 replications) standard errors for Column (4) in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, and a triplet of previous and current residences and workplace: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.I estimate the congestion parameter θ based on the structural relationship between the estimated reduced form parameters ($\beta_R/\beta_G = \theta$): $\hat{\theta} = 0.787\,(0.199)$.

Table 1.4: Semi-Elasticity of Migration with respect to Distance

Dependent Variable:	(1) $\ln \pi_{orm,t}$	(2) $\ln \pi_{orm,t}$	(3) $\ln \pi_{orm,t}$	(4) $\ln \pi_{orm,t}$	(5) $\ln \pi_{or,t}$
distance d_{or} $(-\rho\epsilon)$	-0.002***	-0.004***	-0.033***	-0.033***	-0.007***
	(0.0001)	(0.0001)	(0.0009)	(0.001)	(0.0001)
distance×2005				0.0001	
				(0.002)	
distance×2010				0.0005	
				(0.002)	
Fixed effects:					
Commute Pair ×Year ($\phi_{rm,t}$)	N	Y	Y	Y	N
Job Finding Pair ×Year ($\phi_{om,t}$)	N	N	Y	Y	N
Origin ×Year ($\phi_{o,t}$)	N	N	N	N	Y
Current Residence × Year ($\phi_{r,t}$)	N	N	N	N	Y

Notes: In this table, I estimate the semi-elasticity of migration with respect to distance based on Equation 1.14, starting with a simple estimate without any fixed effects in Column (1) and gradually adding the fixed effects. Column (3) corresponds to Equation 1.14. Column (4) tests whether the semi-elasticity is time-invariant or not. In Column (5), I report the estimated coefficient based on Equation (1.15) following the literature on migration. The sample is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location for Columns (1) - (4). Robust standard errors in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, job finding pair × year, and a triplet of previous and current residences and workplace for Columns (1) - (4) : $***p < 0.01$, $**p < 0.05$, $* < 0.1$. The unit of observation for Column (5) is a pair of previous and current residences. Robust standard errors in parentheses, with three-way clustering by previous residence × year, current residence × year, and migration pair for Column (5): $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

Table 1.5: Semi-Elasticity of Commuting with respect to Distance

	(1)	(2)	(3)	(4)	(5)
Dependent Variable:	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{rm,t}$
distance d_{rm} $(-\kappa\epsilon)$	-0.013***	-0.035***	-0.045***	-0.046***	-0.074***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
distance×2005				0.001	
				(0.001)	
distance×2010				0.003**	
				(0.001)	
Observations	258,323	258,323	258,323	258,323	20,676
Fixed effects:					
Migration Pair ×Year ($\phi_{or,t}$)	N	Y	Y	Y	N
Job Finding Pair ×Year ($\phi_{om,t}$)	N	N	Y	Y	N
Current Residence ×Year ($\phi_{r,t}$)	N	N	N	N	Y
Workplace ×Year ($\phi_{m,t}$)	N	N	N	N	Y

Notes: In this table, I estimate the semi-elasticity of commuting with respect to distance based on Equation (1.16), starting with a simple estimate without any fixed effects in Column (1) and gradually adding the fixed effects. Column (3) corresponds to Equation (1.16). Column (4) tests whether the semi-elasticity is time-invariant or not. In Column (5), I report the estimated coefficient based on Equation (1.17) following the literature on commuting. The sample is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location for Columns (1) - (4). Robust standard errors in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, job finding pair × year, and a triplet of previous and current residences and workplace for Columns (1) - (4): $***p < 0.01$, $**p < 0.05$, $* < 0.1$. The unit of observation for Column (5) is a pair of current residence and workplace location. Robust standard errors in parentheses, with three-way clustering by current residence × year, workplace location × year, and commuting pair for Column (5): $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

Table 1.6: Semi-Elasticity of Job Finding with respect to Distance

	(1)	(2)	(3)	(4)
Dependent Variable:	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$
distance d_{om} (-$\delta\epsilon$)	-0.001***	-0.004***	-0.016***	-0.015***
	(0.00001)	(0.00001)	(0.0003)	(0.0004)
distance×2005				-0.002**
				(0.001)
distance×2010				0.001
				(0.001)
Observations	258,323	258,323	258,323	258,323
Fixed effects:				
Commute Pair ×Year ($\phi_{rm,t}$)	N	Y	Y	Y
Migration Pair ×Year ($\phi_{or,t}$)	N	N	Y	Y

Notes: In this table, I estimate the semi-elasticity of commuting with respect to distance based on Equation (1.18), starting with a simple estimate without any fixed effects in Column (1) and gradually adding the fixed effects. Column (3) corresponds to Equation (1.18). Column (4) tests whether the semi-elasticity is time-invariant or not. The sample is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location. Robust standard errors in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, job finding pair × year, and a triplet of previous and current residences and workplace: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

Table 1.7: Summary of Parameterization

Parameter	Description	Value	Method	Source
α	labor share	0.823	estimated	Economic Census
$1-\beta$	housing expenditure share	0.15	estimated	HH Expenditure Survey
ϵ	shape parameter	3.54	estimated	fixed effects
λ	value of local gov't goods	0.303	estimated	Gravity Equation
θ	net congestion	0.787	estimated	Gravity Equation
ρ	spatial decay of migration	0.009	estimated	Gravity Equation
κ	spatial decay of commuting	0.013	estimated	Gravity Equation
δ	spatial decay of job finding	0.005	estimated	Gravity Equation
τ	income tax rate	0.245	observed	Income Tax Act
ς	local-national revenue sharing	0.091	observed	Local Tax Act
χ	extent of redistribution	0.35	observed	Local Subsidy Act
$\{\varsigma_j\}$	redistribution		observed	Local Subsidy Act
$\{A_j\}$	productivity		recovered	PM+ZP
$\{\tilde{B}_j\}$	adjusted amenities		recovered	fixed effects
$\{H_j\}$	floor space		recovered	Floor space market clearing

Notes: This table summarizes the estimates of the structural parameters of the model. Note that I estimate the value of $1-\beta$ using the Household Expenditure Survey of 2015. Alternatively, based on the estimation result summarized in Table 1.3 and the estimated value of ϵ, I can recover the structural value of $1-\beta$ equal to 0.14. See Section 1.7 for a detailed description of how the parameters above are estimated and recovered.

Table 1.8: Determinants of Redistribution Policy in 2015

	(1)	(2)	(3)	(4)
Dependent Variable	Observed Redistribution Policy ($\ln \varsigma_r$)			
Residential Population ($\ln R_r$)	0.441***	0.530***	0.610***	0.631***
	(0.0269)	(0.0285)	(0.0284)	(0.0578)
Amenities ($\ln \tilde{B}_r$)		-0.170***	-0.184***	-0.187***
		(0.0323)	(0.0255)	(0.0273)
Productivity ($\ln A_r$)			-0.636***	-0.605***
			(0.0906)	(0.130)
Employment Population ($\ln L_r$)				-0.0296
				(0.0730)
Area ($\ln Area_r$)	0.259***	0.262***	0.253***	0.254***
	(0.0146)	(0.0126)	(0.0113)	(0.0111)
Observations	222	222	222	222
R^2	0.619	0.668	0.744	0.744

Notes: In this table, I investigate the determinants of the rules of redistribution by regressing the log of percentage of the total intergovernmental transfers each district receives against local characteristics. The dependent variation is the log of the share of intergovernmental transers each district recieved in 2015. I begin with covariates of residential population and area Column (1) and gradually introduce additional covariates across columns. Each observation correspond to a district in 2015.

1.11 Data Appendix

Wages

I construct wages for each district based on the Economic Census of South Korea in 2015. The Census surveys the universe of establishments in South Korea and records the number of employees and the total costs of labor. I aggregate these two information across establishments in each district and divide the total costs of labor by the number of employees to obtain the district-level wages.

Floor Space Prices

The data source for floor space prices in 2015 is the universe of housing transaction records provided by the Ministry of Land, Infrastructure, and Transport. Each record includes information on the location of a property (district), month and year of purchase, year built, lot size, etc. In order to obtain floor space prices representative for each district in 2015, I employ a Case-Shiller type repeated sales approach at the district level. To to so, I regress log of unit price on a set of dummies for year built, for month of purchase, and for year of purchase excluding 2015 along with district-level fixed effects. I use the estimated values of the district fixed effects (normalized such that the geometric mean is equal to 1) as my data for district-level floor space prices in 2015.

Additional District Level Characteristics

KOSIS (Korean Statistical Information System) provides a wide range of summary statistics describing district-level characteristics. I use the number of firms, number of firms discharging waster water, divorce rates, suicide rates, and geographical land area for each district to carry out cross-validation exercises comparing the model implied values of productivity and amenities with district-level characteristics. In addition, I collected information on the total land area used for residential purposes from the Land Use Statistics publicized by the Ministry of Land, Infrastructure, and Transport.

Annual Migration Rates

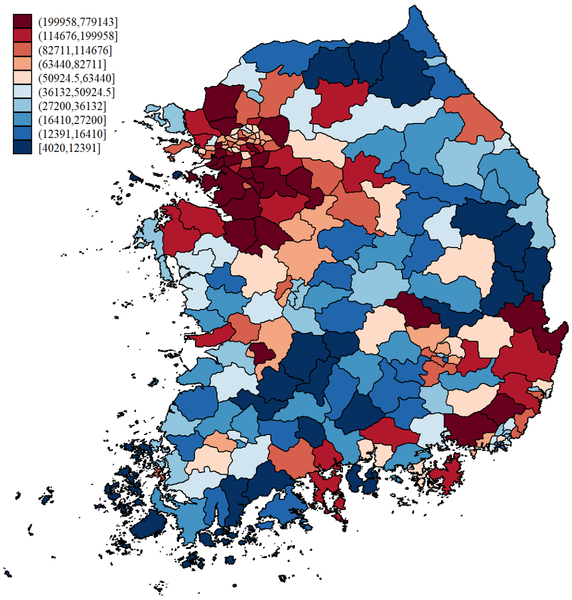
In order to understand the magnitude of migration rates across districts and across provinces (groups of districts), I leverage the restricted-access administrative data, which maintains the universe of migrant registry records in South Korea. This data is not used for the empirical analysis of this paper because the records do not contain where migrants commute to. Notwithstanding its drawback, the records allow me to compute the annual migration rates and compare their magnitudes with the migration rates in the U.S.

1.12 Supplementary Empirical Results

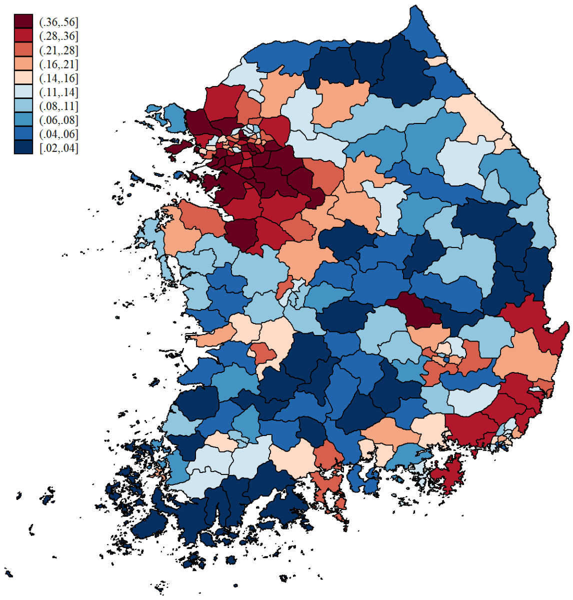
1.12.1 Local Income Taxes and Intergovernmental Transfers

Figure 1.9: Spatial Distribution of Local Government Revenue by Sources

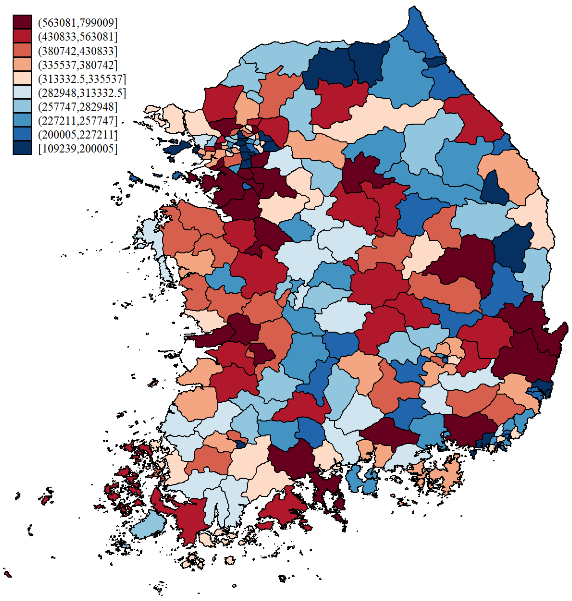
(a) Local Income Tax



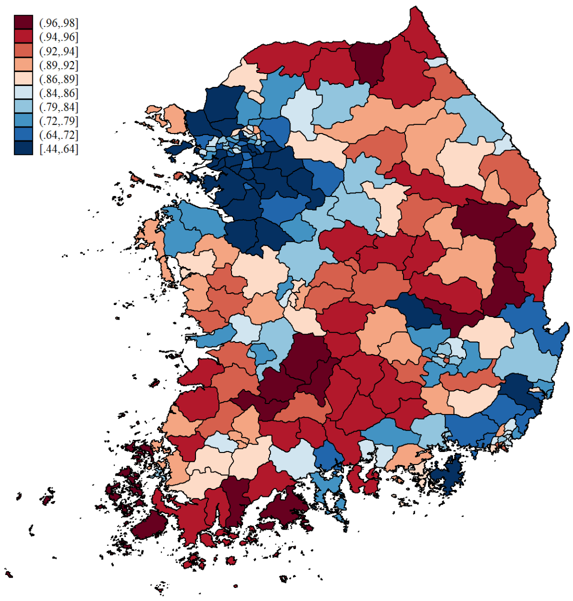
(b) Share of Local Income Tax



(c) Intergovernmental Transfers

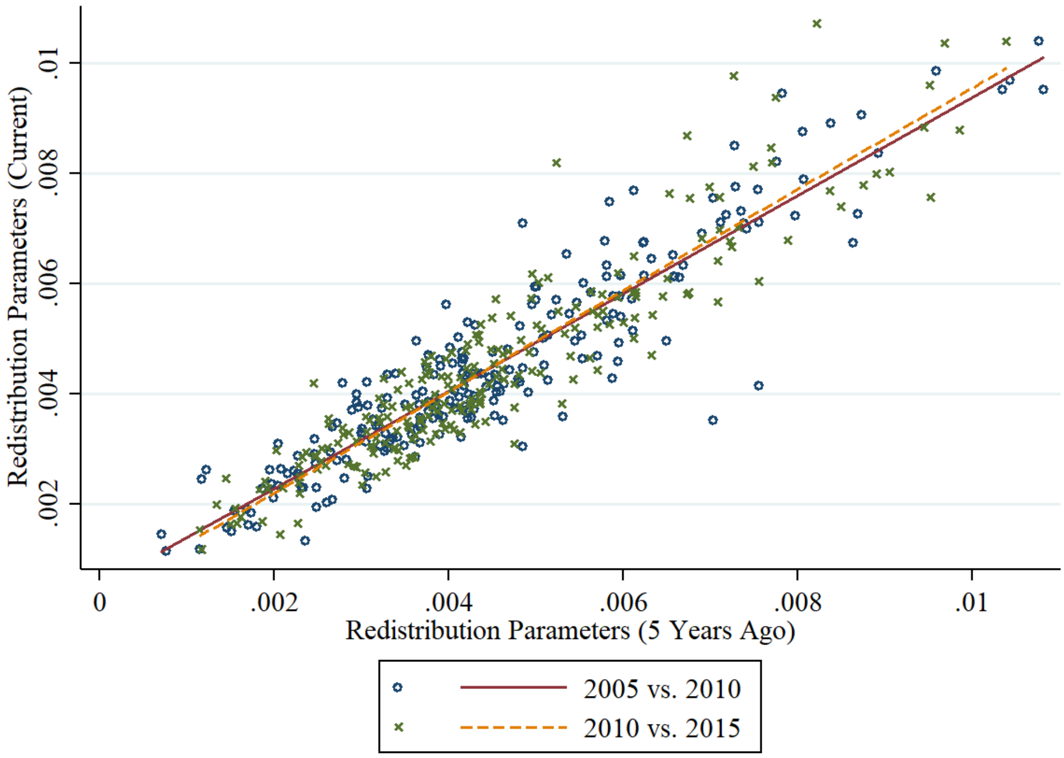


(d) Share of Intergovernmental Transfers



Notes: The figure on the left plots the spatial distribution of local government revenue by its sources (local income taxes and intergovernmental transfers) in 2015. The data source is the administrative data from the Ministry of Interior and Safety of South Korea. The denominator of the shares plotted in Panel (b) and (d) are the sum of local income tax and intergovernmental transfers for each district in 2015.

Figure 1.10: Redistribution Paramters over Time



Notes: This figure plots the shares of the total local tax revenues allocated for intergovermental transfers each locality in a given year again the shares five years ago. The estimated slope is equal to 1 for both (2005 vs. 2010 and 2010 vs. 2015).

1.12.2 Decomposition of Observed Spatial Distribution of Workers

In order to understand the importance of the three spatial linkages (costs of migration, commuting, and job finding) in explaining the observed variation in the spatial distribution of workers $\pi_{orm,t}$, I carry out a type of variance decomposition exercise. First, I purge out location-specific factors $S_{o,t}$, $S_{r,t}$, and $S_{m,t}$ by residualizing $\pi_{orm,t}$ by the location specific fixed effects interacted with year dummies, $\phi_{o,t}$, $\phi_{r,t}$, and $\phi_{m,t}$. Second, I regress the residual on the fixed effects for each bilateral linkage, ϕ_{or}, ϕ_{rm}, and ϕ_{om} and obtain the predicted value $\hat{\pi}_{orm}$.[55] Mechanically, the predicted value is completely explained by ϕ_{or}, ϕ_{rm}, and ϕ_{om} together. By regressing $\hat{\pi}_{orm}$ on each of the pair-wise fixed effects of the

[55] The value of R^2 resulting from this regression is 0.72. Further interacting the set of bilateral fixed effects with year dummes only increases the value of R^2 to 0.79.

bilateral linkages one at a time, I summarize how much of the variation is explained by spatial linkages; the results are summarized in the table below. The migration linkage alone explains 41% of the variation; the commuting linkage explains 8%; the job finding alone explains 10%. Furthermore, an R^2 resulting from accounting any combination of two linkages together is higher than the sum of R^2's resulting from accounting each of the linkages separately.

Table 1.9: Decomposition of Observed Variation in the Data

Regressors	R^2	Regressors	R^2
ϕ_{or}	0.4097	d_{or}	0.0700
ϕ_{rm}	0.0836	d_{rm}	0.0394
ϕ_{om}	0.1000	d_{om}	0.0563
ϕ_{or} and ϕ_{rm}	0.8590	ϕ_{or} and $\phi_{rm} + d_{om}$	0.9051
ϕ_{or} and ϕ_{om}	0.5498	ϕ_{or} and $\phi_{om} + d_{rm}$	0.8729
ϕ_{rm} and ϕ_{om}	0.3216	ϕ_{rm} and $\phi_{om} + d_{or}$	0.5154

Notes: This table reports the values of adjusted R^2 resulting from regressing the predicted spatial distribution of workers $\hat{\pi}_{r_0 rm}$ on the regressors listed in each row. The predicted distribution is computed based on the regression of the observed spatial distribution of workers residualized by location specific factors on the fixed effects of all three bilateral linkages of migration, commuting, and job finding.

Distances between origins and workplace locations together with the fixed effects of the migration and commuting linkages explain 91% of the observed variation. Distances of commuting together with the fixed effects of migration and job finding explain 87%. Lastly, distances of migration with the fixed effects of commuting and job finding accounts for 52% of the observed variations. Based on this simple exercise, I draw the following conclusions. First, net of the location specific factors, the spatial linkages of migration, commuting, and job finding are important determinants of the spatial distribution of workers. Second, the extent to which the commuting linkage explains the variation significantly improves along with the job finding linkage.

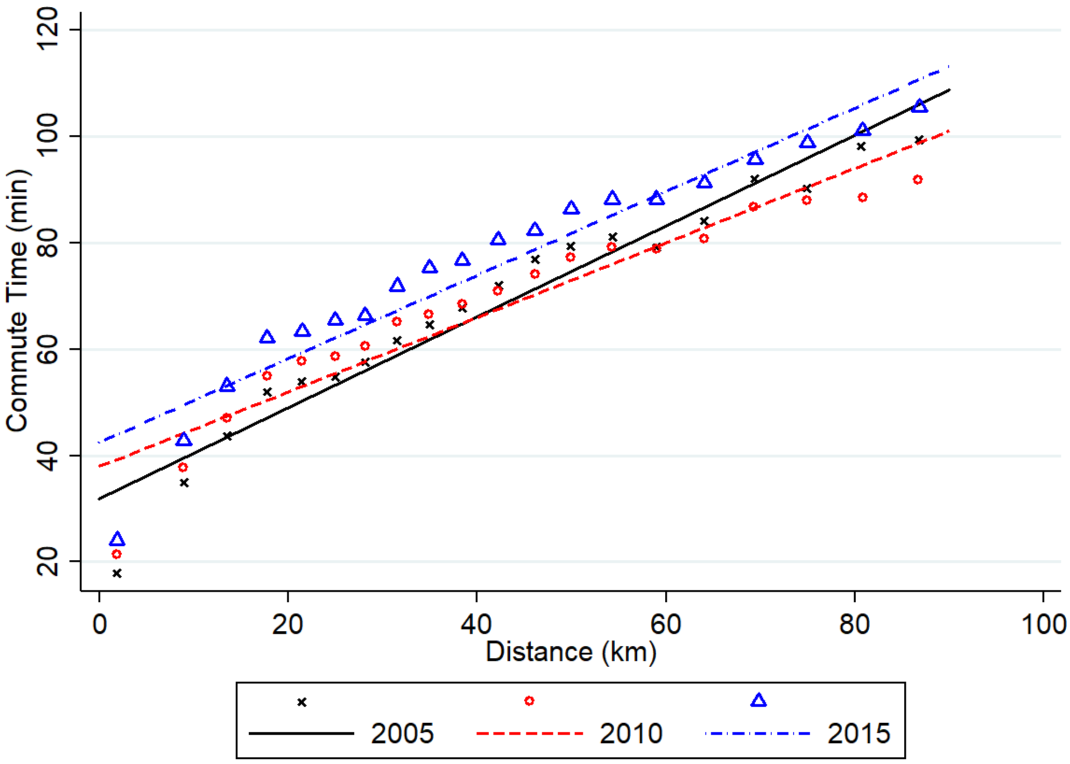
1.12.3 Inference

In this section, I discuss how I address issues related to estimating standard errors estimating the key elasticities of worker mobility. The concern overall is that the errors in each specification can be correlated in two ways. First, there is a classic clustering concern explained in Moulton (1990). Second, one may worry about the serial correlation over time within a panel dimension Bertrand et al. (2004). In order to address these concerns, I report standard errors that are robust to heteroskedasticity and allow multi-way clusterings.

First, with respect to estimating Equation (1.6), I allow errors to correlate across previous residences and across workplace locations sharing the same current residence in a given year. In addition, the serial correlation within each of the panel dimension (a triplet of previous residence, current residence, and workplace location) over time. Second, I conservatively cluster the standard errors at the migration-pair level when estimating Equation (1.14), at the commuting-pair level when estimating Equation (1.16), and at the job-finding-pair level when estimating Equation (1.18).

1.12.4 Travel Time vs. Distance of Commuting

Figure 1.11: Travel Time vs. Distance of Commuting



Notes: This figure plots average commuting time in minutes for each of 5 percentiles of commuting distance for each survey year (2005, 2010, and 2015) of the Population Census of South Korea.

Table 1.10: Commuting Time (min) vs. Distance (km)

	(1)	(2)	(3)	(4)	(5)
Dependent Variable:	Commuting Time ($\tau_{rm,t}^{(time)}$)				
Distance (τ_{rm})	0.788***	0.915***	0.928***	1.017***	
	(0.00801)	(0.00784)	(0.00868)	(0.00839)	
$\tau_{rm} \times 2005$					1.066***
					(0.0121)
$\tau_{rm} \times 2010$					0.919***
					(0.0119)
$\tau_{rm} \times 2015$					1.061***
					(0.0121)
Observations	21,799	21,799	21,799	21,799	21,799
R^2	0.428	0.615	0.598	0.658	0.660
Fixed effects:					
Residence-Year ($\phi_{r,t}$)	N	Y	N	Y	Y
Workplace-Year ($\phi_{m,t}$)	N	N	Y	Y	Y

Notes: This table shows the relationship between distance of commuting and self-reported commuting time reported in the Population Census of South Korea. Each observation is a residence-workplace pair for each year of 2005, 2010, and 2015 with a positive number of workers reported to commute between residential and workplace locations. Robust standard errors in parentheses clustered at the residence-year, the workplace-year level: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

As used in Ahlfeldt et al. (2015) and Morten and Oliveira (2018), an alternative measure to define the cost of commuting is travel time for commuting. Since travel time is surveyed in the Census, I compute average travel times in minutes for all bilateral commuting pairs. Figure 1.11 shows a linear relationship between commuting distance and travel time. Furthermore, inspecting the relationship between commuting distance and time across 2005, 2010, and 2015, there does not seem to be changes in commuting technology. To formalize, I estimate the following specification:

$$time_{rm,t} = \phi_{r,t} + \phi_{m,t} + \kappa^{time} d_{rm} + \varepsilon_{rm,t}^{time}.$$

The results are presented in 1.10. Column (1) shows a raw correlation between distance and time. Across columns, I gradually introduce the fixed effects. According to Column (4), which corresponds to the equation above, travel time of commuting increases when distance of commuting increases by 1 kilometer. In order to understand whether or not this one-to-one relationship is stable over time, I re-estimate the equation above by interacting distance of commuting (time-invariant) with year dummies. The results are summarized in Column (5). The estimated coefficients for 2005, 2010, and 2015 are not statistically different from each other. I conclude that distance is a reasonable proxy for commuting time. The advantage of using travel times may be that measurement errors are averaged out by taking averages of travel times between localities observed at the individual-commuter level. However, average travel time changes over time, and such changes may be correlated with unobserved changes at the residence-workplace pair level that could also affect the spatial distribution of workers (e.g., an introduction of commuter rail). This is not the case for distances as they are fixed over time.

1.12.5 Omitted Variable Bias in OLS Estimates of Elasticities of Worker Mobility with respect to Local Government Goods and Home Prices

Table 1.11: Elasticities of Worker Mobility with respect to Local Government Goods-OVB

	(1)	(2)	(3)	(4)	(5)
Dependent Variable:	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$
$\ln G_{r,t}$ $(\beta_G = \lambda\epsilon)$	-0.231***	0.0965**	-0.433***	-0.452***	0.0957***
	(0.0149)	(0.0405)	(0.0213)	(0.0227)	(0.0299)
$\ln R_{r,t}$ $(\beta_R = \theta\lambda\epsilon)$	0.120***	0.0608**	0.480***	0.482***	0.590***
	(0.0125)	(0.0268)	(0.169)	(0.142)	(0.0522)
$\ln Q_{r,t}$ $(\beta_Q = (1-\beta)\epsilon)$	-0.0416***	-0.0101	-0.0251*	-0.0431***	-0.00148
	(0.0129)	(0.0334)	(0.0135)	(0.0136)	(0.00653)
Observations	258,323	258,323	258,323	258,323	258,323
Fixed Effects:					
Job Finding Pairs $(\phi_{om,t})$	N	Y	N	N	Y
Migration Pairs (ϕ_{or})	N	N	Y	N	Y
Commuting Pairs (ϕ_{rm})	N	N	N	Y	Y

Notes: In this table, I report the OLS estimates of elasticities of worker mobility to local government expenditure, residential density, and home prices based on Equation 1.6, starting with a simple estimate without any fixed effects in Column (1). I introduce the fixed effects for job finding pairs interacted with time in Column (2), the fixed effects for migration pairs in Column (3), and adding the fixed effects for commuting pairs in Column (4). Column (5) reports the OLS estimates with all the fixed effects, separately introduced in Column (2)-(4) and corresponds to Equation 1.6. The sample is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location. Robust standard errors in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, and a triplet of previous and current residences and workplace: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

More formally, the directions of bias with respect to the OLS estimates are expressed as follows:

$$\begin{bmatrix} \widehat{\beta}_G^{OLS} \\ \widehat{\beta}_R^{OLS} \\ \widehat{\beta}_R^{OLS} \end{bmatrix} = \begin{bmatrix} \beta_G \\ \beta_R \\ \beta_Q \end{bmatrix} + \begin{bmatrix} (\sigma_R^2\sigma_Q^2 - \sigma_{RQ}^2)\sigma_{G\varsigma} + (\sigma_{GQ}\sigma_{RQ} - \sigma_{GR}\sigma_Q^2)\sigma_{R\varsigma} + (\sigma_{GR}\sigma_{RQ} - \sigma_{GQ}\sigma_R^2)\sigma_{Q\varsigma} \\ (\sigma_{GQ}\sigma_{RQ} - \sigma_{GR}\sigma_Q^2)\sigma_{G\varsigma} + (\sigma_G^2\sigma_Q^2 - \sigma_{GQ}^2)\sigma_{R\varsigma} + (\sigma_{GR}\sigma_{GQ} - \sigma_{RQ}\sigma_G^2)\sigma_{Q\varsigma} \\ (\sigma_{GR}\sigma_{RQ} - \sigma_{GQ}\sigma_R^2)\sigma_{G\varsigma} + (\sigma_{GR}\sigma_{GQ} - \sigma_{RQ}\sigma_G^2)\sigma_{R\varsigma} + (\sigma_G^2\sigma_R^2 - \sigma_{GR}^2)\sigma_{Q\varsigma} \end{bmatrix} / A \quad (1.38)$$

where $A = \sigma_G^2\sigma_R^2\sigma_Q^2 + 2\sigma_{GR}\sigma_{RQ}\sigma_{GQ} - (\sigma_G^2\sigma_{RQ}^2 + \sigma_R^2\sigma_{GQ}^2 + \sigma_Q^2\sigma_{GR}^2) > 0$ by the Cauchy-Schwarz inequality. Also, note that all the variance and covariance terms are conditional on the set of fixed effects (fixed effects of origin-workplace-by-year, migration pair, commuting pair).

1.12.6 Validity of Instrumental Variables based on the Tax Reforms with Sorting

The quantitative spatial model I present in this paper assumes that the workers are born with initial residences and have heterogeneous preferences for locations. They are otherwise homogeneous. Therefore, I do not take a stance in potential reallocation of workers based on sorting. However, a residence with a greater share of its residents with higher education (skill) may generate a higher amenity value relative to other residences (Diamond, 2016). In this case, the error term in Equation (1.6) would include the distribution of workers by education $\pi_{edu|r,t}$.

The exclusion restriction (1.8) is violated due to sorting only if the fiscal reforms resulted in making residences relatively more or less attractive by changing the educational composition within districts. Since I observe the contemporaneous shares of workers by education levels from the Population Census of South Korea, I can test whether the tax reforms directly affected the educational composition of workers at their residences. I consider the following specification:

$$\pi_{b|r,t} = \phi_r + \eta_{b',b}\tau_{b',t} + \zeta_{b,r,t}, \tag{1.39}$$

where the dependent variable $\pi_{b|r,t}$ is the demeaned fraction of workers with educational level b (low and high, which proxy the low and high income brackets in the tax schedule) living in residence r in year t; the residence fixed effects ϕ_r captures the baseline differences in the dependent variable; $\tau_{b',t}$ is the tax rates in year t for income bracket b'. With the residence fixed effects, if an estimated value of $\eta_{b',b}$ is statistically different from zero, then I reject the hypothesis that the changes in tax rates for income bracket b' had no impact on the changes in the distribution of workers with education level b.

Table 1.12: Tax reforms did not affect education distribution

	(1)	(2)
A. Educational Attainment: Low ($\pi_{low\|r,t}$)		
Tax Rate (Low) $\tau_{low,t}$	3.79e-09	
	(0.00115)	
Tax Rate (High) $\tau_{high,t}$		2.38e-09
		(0.000738)
B. Educational Attainment: High ($\pi_{high\|r,t}$)		
Tax Rate (Low) $\tau_{low,t}$	-1.35e-09	
	(0.000582)	
Tax Rate (High) $\tau_{high,t}$		-6.44e-10
		(0.000382)
Observations	666	666

Notes: This table reports the estimation results based on Equation (1.39). Each estimated coefficient corresponds to the effect of changes in tax rates on changes in the educational composition of residences. The sample is constructed from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,494,198 individual household heads who are employed between the ages of 25 and 60. Each observation corresponds to a residence for each year. Robust standard errors in parentheses clustered at the residence level: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

Table 1.12 reports the estimation results. All the coefficients are not statistically different from zero, nor are they economically significant. In sum, I draw a conclusion that the tax reforms did not result in changes in the attractiveness of residences based on their educational composition of workers. Therefore, predicted tax contributions by low and high income groups are orthogonal to the contemporaneous education distribution.

1.12.7 Alternative (Parsimonious) Specification to Estimate the Elasticities of Worker Mobility with respect to Local Government Goods and Home Prices

$$\begin{aligned}\ln \pi_{orm,t} = \; & \underbrace{\ln \phi_{o,t}}_{\tilde{\phi}_{o,t}} + \underbrace{\ln \phi_{m,t}}_{\tilde{\phi}_{m,t}} + \ln(\varepsilon_{or,t}\varepsilon_{rm,t}\varepsilon_{om,t}\varepsilon'_{orm,t}D_{or}D_{rm}D_{om})^{-\epsilon} \\ & + \beta_G \ln G_{r,t} - \beta_R \ln R_{r,t} - \beta_Q \ln Q_{r,t} + \ln \tilde{B}_{r,t}\end{aligned} \tag{1.40}$$

Equation 1.40 is the expression for the log transformation of the gravity equation (1.5), augmented with the time subscripts wherever applicable. For expository purposes, I unpack the stochastic error term $\varepsilon_{orm,t}$ into four components: $\varepsilon_{orm,t} = \varepsilon_{or,t}\varepsilon_{rm,t}\varepsilon_{om,t}\varepsilon'_{orm,t}$; I assume each of $\varepsilon_{om,t}$ and $\varepsilon'_{orm,t}$ follows a log normal distribution with mean equal to 1. I consider a parsimonious specification alternative to the main estimating equation (1.6) as follows:

$$\ln \pi_{orm,t} = \underbrace{\ln \phi_{o,t}}_{\tilde{\phi}_{o,t}} + \underbrace{\ln \phi_{m,t}}_{\tilde{\phi}_{m,t}} + \phi_{or} + \phi_{rm} + \beta_G \ln G_{r,t} - \beta_R \ln R_{r,t} - \beta_Q \ln Q_{r,t} + \underbrace{\ln \tilde{B}_{r,t} \varepsilon_{om,t} \varepsilon'_{orm,t}}_{\zeta'_{orm,t}} . \tag{1.41}$$

The difference between Equation 1.5 and Equation 1.41 is that the stochastic error term $\varepsilon_{om,t}$ is loaded onto the error term $\zeta'_{orm,t}$ in Equation 1.41. Both specifications are consistent with the model. I summarize the OLS estimates, first-stage estimates, and 2SLS estimates based on Specification 1.41 in Table 1.13. The results are qualitatively and quantitatively similar to the results reported in Table 1.3.

Table 1.13: (2SLS) Parsimonious FE

	(1)	(2)	(3)	(4)	(5)
	OLS	First Stage	First Stage	First Stage	2SLS
Dependent Variable:	$\ln \pi_{orm,t}$	$\ln G_{r,t}$	$\ln R_{r,t}$	$\ln Q_{r,t}$	$\ln \pi_{orm,t}$
Local Government Expenditure, $\ln G_{r,t}$ ($\beta_G = \lambda\epsilon$)	0.106***				1.105***
	(0.016)				(0.190)
Number of Households, $\ln R_{r,t}$ ($-\beta_R = -\theta\lambda\epsilon$)	0.565***				-0.807***
	(0.023)				(0.296)
Floor Space Prices, $\ln Q_{r,t}$ ($-\beta_Q = -(1-\beta)\epsilon$)	-0.0085				-0.528***
	(0.0039)				(0.039)
Predicted Tax Contribution (low), $IV^{low}_{r,t}$		13.82***	7.054***	48.62***	
		(0.210)	(0.152)	(0.901)	
Predicted Tax Contribution (high), $IV^{high}_{r,t}$		13.86***	6.990***	18.86***	
		(0.099)	(0.071)	(0.422)	
Number of Households 30 years ago, $IV^{R}_{r,t}$		0.028***	-0.018***	0.110***	
		(0.001)	(0.001)	(0.005)	
Observations	258,323	258,323	258,323	258,323	258,323
$\hat{\theta}$	-6.164**				0.731**
	(2.176)				(0.152)

Notes: In this table, I compare the OLS estimates and 2SLS estimates of elasticities of worker's mobility to local government expenditure and resident population levels based on Equation 1.41. Column (1) reports the OLS estimates. Column (2) and Column (3) report the first stage results. The 2SLS estimates are reported in Column (4). Across columns, I replace the pairwise fixed effects of job finding interacted with year dummies in the main specification (1.6) with a more parsimonious set of fixed effects for previous residence by year, workplace location by year, and job finding pairs. The sample ($N = 258,323$) is from 3 waves of the Population Census of South Korea in 2005, 2010, and 2015, based on 3,500,232 male household heads who are employed between the ages of 25 and 60. Each observation corresponds to a triplet of previous and current residences and workplace location. Robust standard errors for Column (1), (2), and (3) and bootstrapped (20,000 replications) standard errors for Column (4) in parentheses, with multi-way clustering by migration pair × year, commuting pair × year, and a triplet of previous and current residences and workplace: $***p < 0.01$, $**p < 0.05$, $*p < 0.1$.

1.12.8 2SLS Results based on Migration and Commute Flows

Table 1.14: Estimation Results based on Migration and Commutie Flows

	(1)	(2)	(3)	(4)	(5)	(6)
	Both		Migration Flows		Commuting Flows	
	OLS	2SLS	OLS	2SLS	OLS	2SLS
VARIABLES	$\ln \pi_{orm,t}$	$\ln \pi_{orm,t}$	$\ln \pi_{or,t}$	$\ln \pi_{or,t}$	$\ln \pi_{rm,t}$	$\ln \pi_{rm,t}$
$\ln G_{r,t}$ $(\beta_G = \lambda\epsilon)$	0.0957***	1.072***	0.357***	-1.522***	0.302***	4.935***
	(0.0299)	(0.387)	(0.0197)	(0.188)	(0.0315)	(0.749)
$\ln R_{r,t}$ $(\beta_R = \theta\lambda\epsilon)$	0.590***	-0.844	1.118***	3.205***	1.113***	-3.293***
	(0.0522)	(0.622)	(0.0361)	(0.267)	(0.0512)	(0.852)
$\ln Q_{r,t}$ $(\beta_Q = (1-\beta)\epsilon)$	-0.00148	-0.490***	0.0424***	0.568***	-0.0288***	-2.011***
	(0.00653)	(0.0672)	(0.00386)	(0.0226)	(0.00729)	(0.222)
Observations	258,323	258,323	70,427	70,427	20,676	20,676
Fixed Effects:						
$\phi_{om,t}, \phi_{or}, \phi_{rm}$	Y	Y	N	N	N	N
$\phi_{o,t}, \phi_{or}$	N	N	Y	Y	N	N
$\phi_{m,t}, \phi_{rm}$	N	N	N	N	Y	Y

Notes: In Column (1) and (2), I report the OLS and 2SLS estimates of the effects of local government spending, residential density, and housing prices based on worker mobility defined in terms of both migration and commuting. In Colimn (3) and (4), I report the OLS and 2SLS estimates based on migration flows alone. In Column (5) and (6), I report the OLS and 2SLS estimates on commuting flows alone.

Recall the gravity equation of the model is given by:

$$\pi_{orm,t} = \frac{(\tilde{B}_{r,t}(1-\tau_{m,t})\tilde{w}_{m,t}G^{\lambda}_{r,t})^{\epsilon}\pi_{o,t}}{(\varepsilon_{orm,t}D_{or}D_{rm}D_{om}Q^{1-\beta}_{r,t}R^{\theta\lambda}_{r,t})^{\epsilon}} \Big/ \underbrace{\sum_{r'=1}^{J}\sum_{m'=1}^{J} \frac{(\tilde{B}_{r',t}(1-\tau_{m',t})\tilde{w}_{m',t}G^{\lambda}_{r',t})^{\epsilon}}{(\varepsilon_{or'm',t}D_{or'}D_{r'm'}D_{om}Q^{1-\beta}_{r',t}R^{\theta\lambda}_{r',t})^{\epsilon}}}_{\Phi_{o,t}}$$

Summing it over workplace location, I derive an expression for migration flow:

$$\pi_{om,t} = \frac{(\tilde{B}_{r,t}G^{\lambda}_{r,t})^{\epsilon}\pi_{o,t}/\Phi_{o,t}}{(D_{or}Q^{1-\beta}_{r,t}R^{\theta\lambda}_{r,t})^{\epsilon}} \underbrace{\sum_{m=1}^{J}\left(\frac{(1-\tau_{m,t})\tilde{w}_{m,t}}{\varepsilon_{orm,t}D_{rm}D_{om}}\right)^{\epsilon}}_{ALMA_{or,t}}$$

Then, I derive an estimating equation by taking the log transformation:

$$\ln \pi_{or,t} = \phi_{o,t} + \phi_{or} + \underbrace{\lambda\epsilon}_{\beta_G} \ln G_{r,t} - \underbrace{\theta\lambda\epsilon}_{\beta_R} \ln R_{r,t} - \underbrace{(1-\beta)\epsilon}_{\beta_Q} \ln Q_{r,t} + \zeta^{mig}_{or,t}, \tag{1.42}$$

where $\phi_{o,t} = \ln \phi_{o,t}/\Phi_{o,t}$; $\phi_{or} = \ln D^{\epsilon}_{or}$; $\zeta^{mig}_{or,t} = \ln \tilde{B}_{r,t} ALMA_{or,t}$. The relevance of the instrumental variables (tax reforms and historical residential density) holds as when using the worker mobility (migration and commuting jointly). The exclusion restriction requires:

$$E\begin{bmatrix} IV^{low}_{r,t}\zeta^{mig}_{or,t} & |\phi_{o,t},\, \phi_{or} \\ IV^{high}_{r,t}\zeta^{mig}_{or,t} & |\phi_{o,t},\, \phi_{or} \\ IV^{R}_{r,t}\zeta^{mig}_{or,t} & |\phi_{o,t},\, \phi_{or} \end{bmatrix} = 0.$$

This is violated because $IV^{low}_{r,t}$ and $IV^{high}_{r,t}$ are functions of tax rates. And, the error terms includes $ALMA_{or,t}$, which is also a function of tax rates. Therefore, 2SLS estimates would be inconsistent. In particular, the direction of bias in 2SLS estimate for β_G is downward $\because$ $cov(G_{r,t}, \zeta^{mig}_{or,t} < 0)$. Note that if D_{rm} is equal to 1 (i.e., no spatial friction from commuting), then exclusion restriction is satisfied. Therefore, observing biases in 2SLS estimates using migration implies it is important to take commuting into account.

Similarly, I sum $\pi_{orm,t}$ over previous residence, I derive an expression for commuting flow:

$$\pi_{rm,t} = \frac{(\tilde{B}_{r,t}(1-\tau_{m,t})\tilde{w}_{m,t}G^{\lambda}_{r,t})^{\epsilon}}{(D_{rm}Q^{1-\beta}_{r,t}R^{\theta\lambda}_{r,t})^{\epsilon}} \underbrace{\sum_{o=1}^{J} \frac{\pi_{o,t}/\Phi_{o,t}}{(\varepsilon_{orm,t}D_{or}D_{om})^{\epsilon}}}_{AMMA_{rm,t}}$$

Then, I derive an estimating equation by taking the log transformation as follows:

$$\ln \pi_{rm,t} = \phi_{m,t} + \phi_{rm} + \underbrace{\lambda\epsilon}_{\beta_G} \ln G_{r,t} - \underbrace{\theta\lambda\epsilon}_{\beta_R} \ln R_{r,t} - \underbrace{(1-\beta)\epsilon}_{\beta_Q} \ln Q_{r,t} + \zeta^{com}_{rm,t}, \qquad (1.43)$$

where $\phi_{m,t} = \ln((1-\tau_{m,t})\tilde{w}_{m,t})^{\epsilon}$;$\phi_{rm} = \ln D^{\epsilon}_{rm}$; $\zeta^{com}_{rm,t} = \ln \tilde{B}_{r,t} AMMA_{rm,t}$. Similarly, the relevance of the instrumental variables (tax reforms and historical residential density) still holds since the relevance did not hinge on the assumptions of spatial frictions. The exclusion restriction requires,

$$E\begin{bmatrix} IV^{low}_{r,t}\zeta^{com}_{or,t} & |\phi_{m,t},\, \phi_{rm} \\ IV^{high}_{r,t}\zeta^{com}_{or,t} & |\phi_{m,t},\, \phi_{rm} \\ IV^{R}_{r,t}\zeta^{com}_{or,t} & |\phi_{m,t},\, \phi_{rm} \end{bmatrix} = 0.$$

The exclusion restriction in this case is violated. In this case, $IV^{low}_{r,t}$ and $IV^{high}_{r,t}$ are functions of tax rate, while the error terms includes $AMMA_{or,t}$ also a function of tax rates because $\Phi_{o,t}$ includes tax rates. Therefore, 2SLS estimates would be inconsistent.

In particular, the direction of bias in 2SLS estimate for β_G is upward $\because cov(G_{r,t}, \zeta^{com}_{rm,t} > 0)$. Note that if D_{or} is equal to 1 (no spatial friction from migration), then exclusion restriction is satisfied. Therefore, observing biases in 2SLS estimates based on commuting flows implies that migration needs to be taken into account.

In Table 1.14, I report the OLS and 2SLS estimates based on Equation (1.6) in Column (1) and (2) using both migration and commuting flows, Equation (1.42) in Column (3) and (4) using migration flows alone, and Equation (1.43) using commuting flows alone. The results altogether show that in order to consistently estimate the elasticities of interest leveraging the tax reforms, both migration and commuting need to be considered jointly.

($\phi_{m,t}$-OVB) If a district is located near employment locations with high wages, then the average income of the residents in this district is high. Because of the redistributive intergovernmental transfers, the local government expenditure is low. The fixed effects for employment locations address the omitted variable bias rising from the negative correlation between local labor market returns and local government spending.

($\phi_{m,t}$-Exclusion Restriction) Without the fixed effects for workplace location, the exclusion restriction of the proposed instruments based on tax reforms is violated because the tax rates directly affects worker mobility.

($\phi_{o,t}$-OVB) There are two factors specific to origin: the initial distribution of workers across residential location and multilateral resistance. So, the fixed effects capture the effects of augmented migrant market access. If a residence is situated around places with higher values of migrant market access, this residence is likely to have a greater number of migrants, resulting in a higher residential density. In turn, higher population is likely to be positively correlated with local spending because of a high tax base and redistribution favoring dense localities.

($\phi_{o,t}$-Exclusion Restriction) Exclusion redistribution is violated unless these fixed effects are introduced because the multilateral resistance term is a function of tax rates.

1.13 Supplementary Quantitative Results

1.13.1 Adjusted After-Tax Wages and Fréchet Shape Parameter

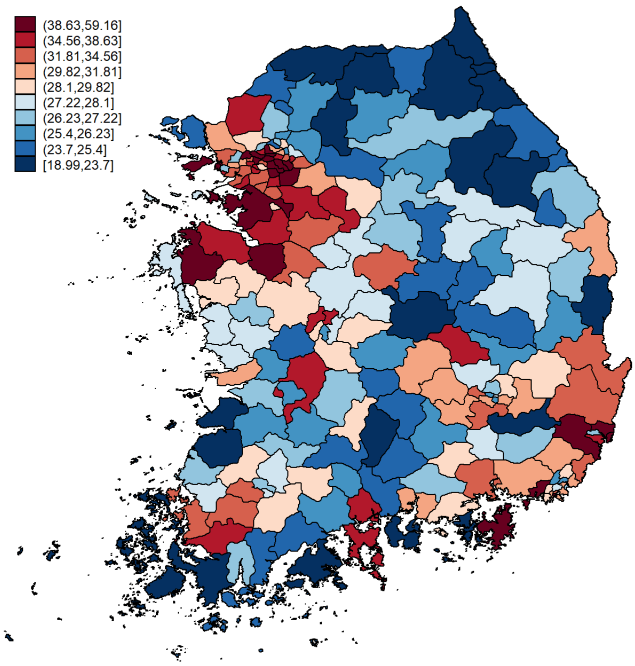
$$\pi_{m|rr_0} = \frac{M_m \left(\frac{(1-\tau_m)w_m}{D_{r_0m}D_{rm}}\right)^{\epsilon}}{\sum_{m'=1}^{S} M_{m'} \left(\frac{(1-\tau_{m'})w_{m'}}{D_{r_0m'}D_{rm'}}\right)^{\epsilon}} \tag{1.44}$$

Take log transformation both sides. Then, add the costs of job finding and commuting. Then, regress the left hand side variable on the fixed effects of workplace location and migration pairs. I recover the adjusted after-tax wages from the estimated fixed effects of workplace locations.

I estimate the dispersion parameter by taking the ratio of the dispersion of adjusted after-tax wages and the dispersion of observed after-tax wages. The estimated value of ϵ is equal to 3.5, which is statistically significantly different from zero at the 1 percent.

1.13.2 Local Productivity

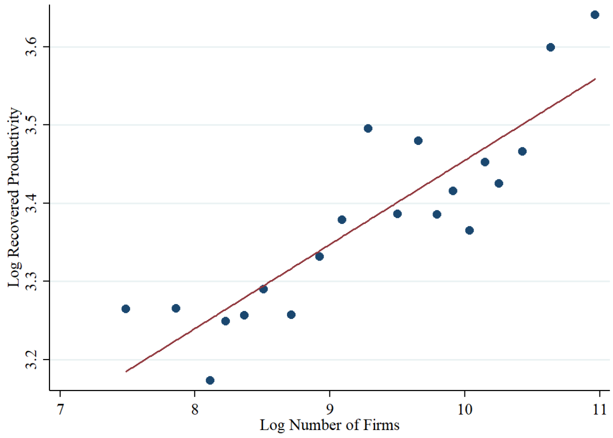
Figure 1.12: Spatial Distribution of Productivities



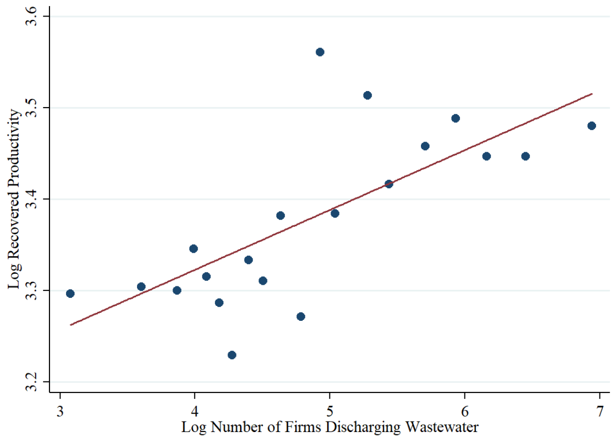
Notes: This figure plots the recovered values of productivity for each district using the model with the data in 2015. Section 1.7.3 explains how the values are recovered from the estimated fixed effects.

Figure 1.13: Recovered Local Amenities vs. Number of Firms

(a) Number of Firms



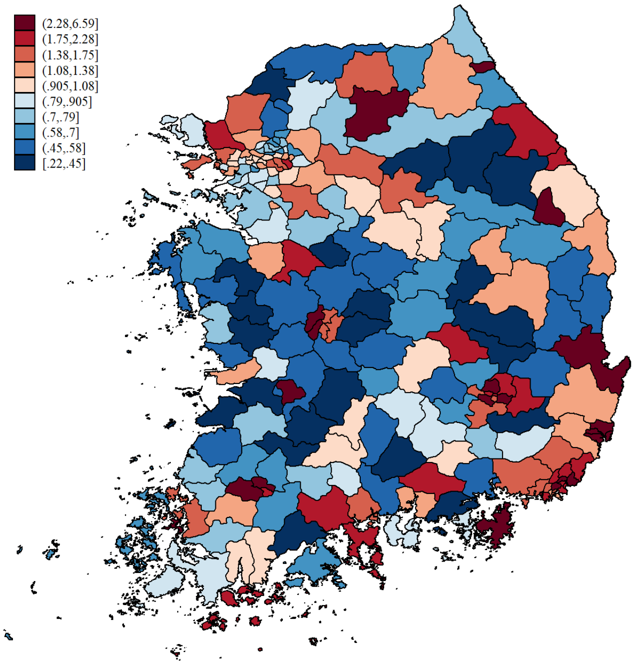
(b) Number of Dirty Firms



Notes: This figure plots the values of log productivity recovered in Section 1.7.3 against number of firms in Panel (a) and firms discharging wastewater in Panel (b) in 2015. Each point corresponds to 5 percentile of number of firms.

1.13.3 Adjusted Amenities

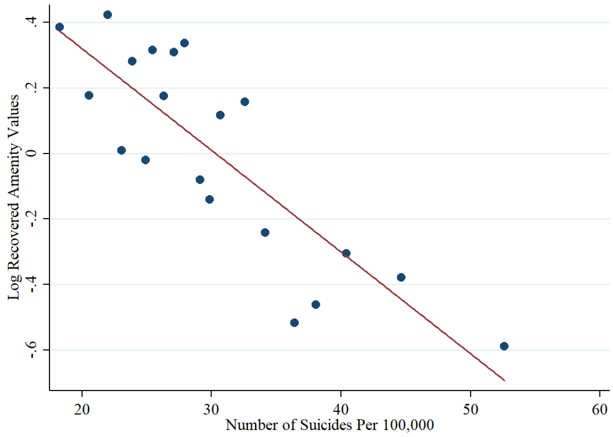
Figure 1.14: Recovered Amenities in 2015



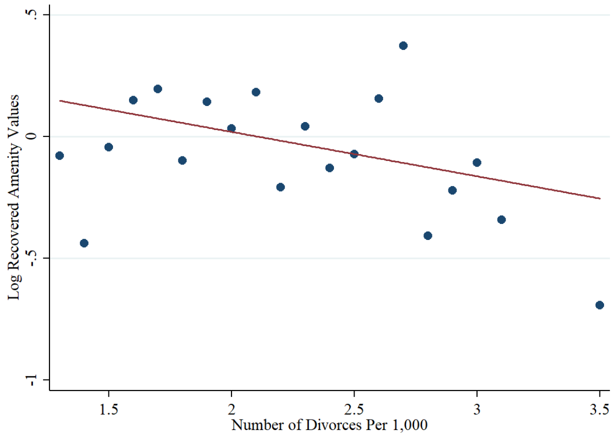
Notes: This figure plots the recovered amenity values for each district using the model with the data in 2015. Section 1.7.3 explains how the amenity values are recovered from the estimated fixed effects.

Figure 1.15: Recovered Local Amenities vs. Measures of Quality of Life

(a) Suicide Rate



(b) Divorce Rate



Notes: This figure plots the values of adjusted amenities recovered in Section 1.7.3 against two measures proxying the quality of life observed in 2015: suicides per 100,000 residents in Panel (a); number of divorces per 1,000 couple in Panel (b). Each point corresponds to 5 percentile of the quality-of-life measures.

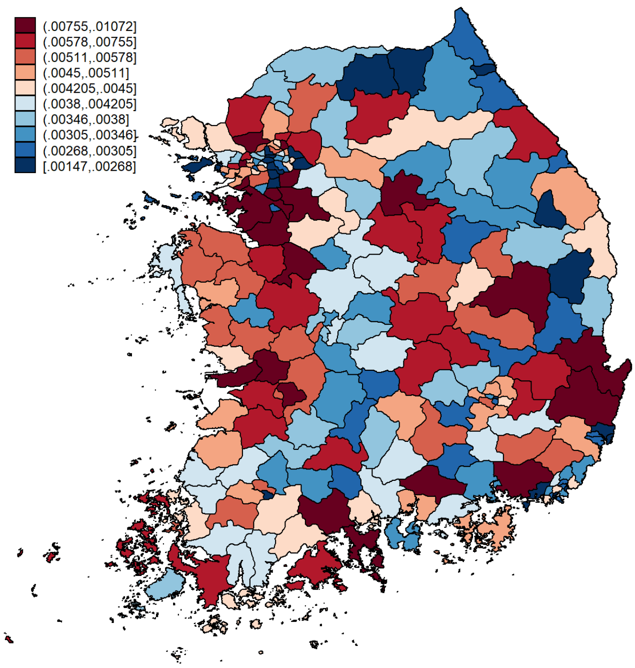
1.13.4 Fiscal Decentralization Policy Parameters

- Observed Data: Total Expenditure G_r and its sources: local tax revenue LT_r and intergovernmental transfers IT_r
- Local government spending:

$$G_r = \varsigma \sum_{m=1}^{S} \tau_m w_m \pi_{m|r} R_r + \varsigma_r (1-\varsigma)\chi \sum_{r'=1}^{S} \sum_{m'=1}^{S} \tau_{m'} w_{m'} \pi_{m'|r'} R_{r'} \tag{1.45}$$

where $\tilde{\varsigma}$ denotes the fraction of total local tax revenue delivered to the national government $1-\varsigma = 0.9$ multiplied by the fraction of the national tax revenue used for redistribution $\chi = 0.35$. This implies that $(1-\chi)(1-\varsigma) = 0.5915$ of the local tax revenue is used for the national government. This also means that in total about 40 percent of the local tax revenue (i.e., extent of fiscal decentralization) is spendt locally. When I conduct counterfactual policy experiments. I keep the extent of fiscal decentralization constant at 40 percent and only change the extent of redistribution, ranging from 0 to 40 percent. Also, I keep the rules of redistribution ($\{\varsigma_j\}_{j=1}^{J}$, where $\varsigma_j = \frac{IT_j}{\sum_{j'=1}^{J} IT_j}$) constant at the 2015 values. When the extent of redistribution is equal to 0%, local government spending is solely financed by local tax revenue from residents. When it is equal to 40%, local government spending is completely determined by intergovernmental transfers.

Figure 1.16: Redistribution Parameters in 2015



Notes: This figure plots the observed values of rules of redistribution (ς_r) in 2015 for each residence. The districts in red receives greater shares of intergovernmental transfers from the national government than the districts in blue. The shares sum up to 1.

1.13.5 Algorithm to Solve the Model

I briefly describe the iterative algorithm used to solve for the equilibrium of the model (Ahlfeldt et al., 2015; Monte et al., 2018; Tsivanidis, 2019); See Appendix 1.13.5 for details. Section 1.6.5 characterizes the equilibrium of the model and the system of equations to be solved. First, I make initial guess for a set of endogenous variables. Second, using these initial values, I solve the system of equations of the model for a new value of the endogenous variables. Third, I update the guess for the equilibrium by taking a weighted average of the initial and the new values. Lastly, I iterate this process until the new and initial values converge.

I solve for stock of floor space for each district appealing to the market clearing for floor space in Section 1.6.3. First, in equilibrium, the residential floor space demanded is a function of after-tax wages ($(1-\tau)w_m$), conditional commuting probabilities ($\pi_{m|r}$), residential population (R_r), and per-unit floor space prices (Q_r) given housing expenditure share ($1-\beta$) as in Equation (1.27). Second, the commercial floor space demanded is determined by local productivity (A_j), employment population (L_j), and floor space prices (Q_j) given labor share in production (α) as in Equation (1.28). I set floor space stock of a district equal to the sum of floor space demands for residential and commercial uses computed based on the tax rate from Section 1.7.2 and local productivity recovered above as well as the observed data on wages, floor space prices, conditional commuting probabilities, and residential and employment population.

1.13.6 Goodness of Fits relative to Alternative Specifications

Table 1.15: Goodness of Fits: Model vs. Alternatives (Migration)

	Restrictions					
	(1)	(2)	(3)	(4)	(5)	(6)
$\frac{\sum(X_r^{restriction}-X_r^{data})^2}{\sum(X_r^{baseline}-X_r^{data})^2}$	$\rho=0$	$\delta=0$	$\rho=0$ $\delta=0$	$\rho=0$ $\delta=0$ $\kappa=\kappa^{lit}$	$\kappa\to\infty$	$\rho=\rho^{lit}$ $\delta=0$ $\kappa\to\infty$
$X_r=$						
R_r	1.16	1.09	2.47	2.61	1.34	3.60
G_r	1.12	1.10	1.87	1.86	1.95	2.83
L_m	1.19	1.14	3.16	3.17	1.97	4.81

Notes: In this table, I show the goodness of fits under alternative assumptions on the spatial frictions relative to the baseline model. Each value reported in this table corresponds to the sum of squared residuals relative to the baseline model. Therefore, a value higher than 1 implies that the base line model performs better in predicting the observed values. I solve for an equilibrium assuming that migration is costless in Column (1), job finding is costless in Column (2), both are costless in Column (3), and both are costless with the spatial decay of commuting equal to the value estimated following the commuting literature (Column (5) of Table 1.5). In Column (5), I assume that commuting is prohibitively costly. Column (6) assumes the value of spatial decay of migration equal to the value estimated following the migration literature (Column (5) of Table 1.4) while job finding is costless.

1.14 Supplementary Theoretical Results

1.14.1 Derivation of the Gravity Equation in Section 1.3

Because the indirect utility is equal to the idiosyncratic component of utility (z_{irm}) multiplied by the indirect utility of the systematic component (v_{orm} in Equation (1.3)), the distribution of utility for a worker from origin o living in district r and working in district m is also Fréchet distributed. Therefore, the cumulative distribution function of the utility is

$$F_{rm}(u) = \Pr[U \leq u] = \Pr(z \leq u \times v_{orm}^{-1}), \tag{1.46}$$

where $z \sim G(z) = \exp(-T_r M_m z^{-\epsilon})$. It follows that

$$F_{rm}(u) = \exp(-\frac{T_r M_m B_r(1-\tau_m)w_m}{D_{orm}Q_r^{1-\beta}}\left(\frac{G_r}{R_r^{\theta}}\right)^{\lambda} u^{-\epsilon}) \equiv \exp(-\Phi_{orm}u^{-\epsilon}). \tag{1.47}$$

I denote f_{rm} to be the density function. Conditional on their origin o, workers choose a pair of residence r and workplace m that achieves that maximum utility. Therefore, the probability of choosing a residence-workplace pair (residence r and workplace location m) conditional on having come from origin o is expressed as follows:

$$\begin{aligned}
\pi_{rm|o} &= \Pr[u_{rm|o} \geq \max\{u_{jk}\}; \forall j, k] \\
&= \int_0^{\infty} \prod_{k \neq j} F_{rk}(u) \times \left(\prod_{j \neq r}\prod_{k} F_{jk}(u)\right) f_{rm}(u)du \\
&= \int_0^{\infty} \prod_{j}\prod_{k} \epsilon\Phi_{orm}u^{-(\epsilon+1)}\exp(-\Phi_{ojk}u^{-\epsilon})du \\
&= \int_0^{\infty} \epsilon\Phi_{orm}u^{-(\epsilon+1)}\exp(-\Phi_{o}u^{-\epsilon})du,
\end{aligned}$$

where $\Phi_o = \sum_{r=1}^{J}\sum_{m=1}^{J}\Phi_{orm}$. Evaluating the integral above, the probability of choosing residence r and workplace m conditional on origin o is:

$$\pi_{rm|o} = \frac{T_r M_m \left(\frac{B_r(1-\tau_m)w_m}{D_{orm}Q_r^{1-\beta}}\left(\frac{G_r}{R_r^{\theta}}\right)^{\lambda}\right)^{\epsilon}}{\sum_{r'=1}^{J}\sum_{m'=1}^{J} T_{r'}M_{m'}\left(\frac{B_{r'}(1-\tau_{m'})w_{m'}}{D_{or'm'}Q_{r'}^{1-\beta}}\left(\frac{G_{r'}}{R_{r'}^{\theta}}\right)^{\lambda}\right)^{\epsilon}} \equiv \frac{\Phi_{orm}}{\Phi_o} \tag{1.48}$$

Because the maximum of a sequence of Fréchet distributed random variables is itself Fréchet distributed. Therefore,

$$F_o(u) = \exp(-\Phi_o u^{-\epsilon}), \text{ where } \Phi_o = \sum_{r=1}^{J}\sum_{m=1}^{J}\frac{T_r M_m B_r(1-\tau_m)w_m}{D_{orm}Q_r^{1-\beta}}\left(\frac{G_r}{R_r^{\theta}}\right)^{\lambda}. \tag{1.49}$$

Based on the distribution of utility defined above, the expected utility for workers with origin o is given by:

$$\mathbb{E}[u|o] = \int_0^\infty \epsilon \Phi_o u^{-\epsilon} e^{-\Phi_o u^{-\epsilon}} du = \Gamma\left(\frac{\epsilon-1}{\epsilon}\right) \Phi_o^{1/\epsilon} \equiv \bar{u}_o. \tag{1.50}$$

1.14.2 Isomorphism of the Gravity Equation

I show that the gravity equation (1.5) is isomorphic to the types of gravity equations derived in the literature on costly movements of people: commuting and migration.

1.14.2.1 Commuting Literature

The literature on commuting decisions assume free mobility in terms of migration. Therefore, there is usually no discussion on how workers are distributed across space before they make their commuting decisions. The underlying assumption in this literature is that there is no cost of enabling each commuting possibility via migration and job finding. This assumption translate to setting both ρ and δ equal to zero in my model presented in Section 1.3. Then, the distribution of workers by current residence and workplace is independent to the distribution of workers by initial residence. Therefore, Equation (1.5) does not vary by initial residence o and is given by:

$$\pi_{rm} = \frac{T_r M_m \left(\frac{B_r(1-\tau_m)w_m}{D_{rm} Q_r^{1-\beta}} \left(\frac{G_r}{R_r^\theta}\right)^\lambda\right)^\epsilon}{\sum_{r'=1}^S \sum_{m'=1}^S T_{r'} M_{m'} \left(\frac{B_{r'}(1-\tau_{m'})w_{m'}}{D_{r'm'} Q_{r'}^{1-\beta}} \left(\frac{G_{r'}}{R_{r'}^\theta}\right)^\lambda\right)^\epsilon}, \tag{1.51}$$

where D_{rm} is a commuting cost, a function increasing in distance between r and m. Further assuming no tax on wage (i.e., $\tau_m = 0$ for all m) and no utility derived from local government goods and services (i.e., $\lambda = 0$), Equation (1.51) is identical to the gravity equations based on the spatial models of Ahlfeldt et al. (2015) and Monte et al. (2018).

1.14.2.2 Migration Literature

The literature on migration decisions generally considers movements of people across relatively larger spatial units such that workers are likely to work and live in the same spatial unit upon migrating. Accordingly, in this literature, there is no distinction between a workplace and a residence since workers are assumed to work and live in the same locations. This assumption can be implemented in my model by setting the commuting cost to a workplace outside of residence equal to ∞. Then, the migration patterns of workers are summarized by:

$$\pi_{or} = \frac{T_r M_r \left(\frac{B_r (1-\tau_r) w_r}{D_{or} Q_r^{1-\beta}} \left(\frac{G_r}{R_r^{\theta}} \right)^{\lambda} \right)^{\epsilon} \pi_o}{\sum_{r'=1}^{S} T_{r'} M_{r'} \left(\frac{B_{r'} (1-\tau_{r'}) w_{r'}}{D_{or'} Q_{r'}^{1-\beta}} \left(\frac{G_{r'}}{R_{r'}^{\theta}} \right)^{\lambda} \right)^{\epsilon}}, \tag{1.52}$$

where D_{or} is the iceberg cost associated with migration. Again, assuming to tax on wage and no benefits from local government goods and services, Equation (1.52) shares the same structure as the gravity equations based on the spatial models of migration considered in Bryan and Morten (2018) and Morten and Oliveira (2018).

Chapter 2

Do Pro-Natalist Cash Transfers Work? Evidence from Local Programs in South Korea

Wookun Kim, UCLA[1]

Many countries now experience fertility rates below the 2.1-replacement level of fertility and have experimented with some pro-natalist policy measures to encourage childbearing. Since 1983, the total fertility rate of South Korea has stayed below the replacement level, reaching the lowest in the world in 2005. This paper exploits a unique setting in South Korea to identify the causal effects of pro-natalist cash transfers on birth outcomes. In particular, a cash transfer of 1,000 USD increased the total fertility rate by 0.022 children per woman or 1.8 percent. Next, decomposing total fertility rates by birth parity and age of mothers reveals that the pro-natalist cash transfers had parity-specific effects and did not have effects on fertility rates of adolescents. Lastly, I find no evidence of changes in health outcomes at birth and son preference due to the cash transfers.

[1] Link to most recent version: www.wookunkim.com/research. I thank Adriana Lleras-Muney, Kathleen McGarry, Youssef Benzarti, and Manisha Shah for their encouragement and guidance. I also thank many participants at UCLA applied-micro seminars and the annual PAA meeting in Denver 2018 for helpful comments. All errors are mine.

2.1 Introduction

Many countries have undergone fertility transitions, shifting from 5 to 8 children per woman as recently as 1980 to 2 or fewer children per woman (Strulik et. al., 2015). In many countries, fertility rates are below the replacement level of 2.1 children per woman. Policy makers have expressed growing concerns about the negative impacts of low fertility rates, further exacerbated by an aging population (Morgan, 2003; Frejka et. al., 2010; Harper, 2014). Many developed countries now have implemented policies to boost fertility such as cash transfers, paid and unpaid parental leave for childbearing, and tax benefits.[2]

In this paper, I investigate the effects of local pro-natalist programs in South Korea on fertility. In South Korea, local governments provided congratulatory cash transfers to parents for having babies to rectify national fertility declines. The generosity of cash transfers varied widely across localities, birth orders or parities. To identify the causal effects of local pro-natalist programs on fertility outcomes, I exploit this rich variation in cash transfer amounts and policy implementation timing across 222 districts of South Korea over a period of 15 years. The fertility outcomes I study include both quantity of births (i.e., fertility rates and parity-specific birth rates) and quality of health outcomes (i.e., birth weight and gestation).[3] I collected local policy information from four different sources, and constructed panel data by merging the local pro-natalist cash transfer, with local characteristics, including fertility outcomes such as total fertility rates, age-specific birth rates, and age-specific parity-specific birth rates.

The results suggest a positive and significant effect of pro-natalist cash transfer on the number of births, but no meaningful effects on birth weight and pregnancy duration. First, total fertility rates increased by 0.0216 children per woman, or 2 percent, with a cash transfer of 1,000 USD. Second, when decomposing the effects of cash transfer by birth parity, I find that cash transfer for a birth parity affected the birth rates of the corresponding parity, but did not change the birth rates of the other birth parities. In other words, the effects of cash transfer are parity-specific. Third, the cash transfer only increased the fertility rates between the ages of 20 and 39 of the female population, more active in making fertility decisions, and had no impact on the fertility rate of adolescents or females older than 40. Fourth, by studying the potential policy effects on the female marriage age and average age of mothers at birth, I conclude that the positive effects of

[2] Fleckenstein and Lee (2014) provide detailed pro-natalist policy changes in Britain, Germany, South Korea, and Sweden; Frejka et. al. (2010) summarize pro-natalist policies implemented in East Asia. Gauthier (2007) presents a literature review of findings on the effects of various pro-natalist policies.

[3] During the sample period, some districts were merged and split. I restrict the sample to 222 districts, which did not undergo redistricting, and construct a balanced panel of districts.

cash transfer on the number of births is driven by an increase in the number of children ever born per woman, rather than changes in birth timing. I also find no changes in pregnancy duration and birth weight, implying no meaningful changes in the qualitative aspects of fertility from the cash transfer. Lastly, given the prevalent son preference in the context of South Korea (Edlund and Lee, 2013), I test the policy effects on sex composition at birth and find no evidence of strengthening of son preference.

The key identification assumption is that, conditional on covariates and given province-by-year fixed effects, the generosity of cash transfer across districts is orthogonal to quantity and quality of birth. A potential threat to identification arises if the cash-transfer generosity is correlated with unobserved district characteristics that are also correlated with fertility outcomes. To address this concern, first I show that the results are robust to changing the course of identifying variation from cross-sectional to over-time within each district. Furthermore, I implement an instrumental variable strategy and show that the 2SLS estimator produces similar results to the results based on the fixed effects model. Lastly, I conduct a placebo test by permuting histories of cash transfer across districts to supplement the consistency of the results.

My contribution to the literature is two-fold. First, to the best of my knowledge, this is the first paper to estimate causal effects of pro-natalist cash transfer on birth rates by taking advantage of variation in cash transfer generosity across time and space. As a consequence of data limitations, the previous literature on cash transfer for childbearing has mainly focused on difference-in-difference strategies and regression discontinuity designs, and comparing the fertility outcomes before and after cash transfer implementation, or a one-time change in generosity in a region, while using unaffected regions or ineligible families as a control group (Milligan, 2005; Boccuzzo et. al., 2008; Cohen et. al., 2013; González, 2013). Hong et. al. (2016) have also examined the local government transfers in South Korea. Their results may suffer from errors-in-variables bias and weak moment condition problem (Newey and Windmeijer, 2009). I consider this paper a complement to their with a improved policy data set, alternative identification strategies, and new findings on the heterogeneous effects of the baby bonus across birth parities and ages of mothers. My analysis shows that the cash transfers had parity-specific effects: cash transfers for a birth parity only affect the birth rates of that birth parity, not those of the lower or higher birth order. Furthermore, I provide evidence that the changes in the fertility rate from the cash transfers are driven by an increase in completed fertility, as opposed to changes in the timing of childbearing.

Second, I examine the effects of pro-natalist cash transfer on the qualitative dimension of fertility in addition to the quantity of births. Given that the explicit objective of pro-natalist policies is to increase the number of births, it is important to test if an increase in births brings about decrease in the quality of those births. There is limited evidence on whether or not pro-natalist cash transfer affects the quality of birth. Amarante et. al. (2016) find that cash transfer reduced the incidence of low birthweight for poor families in Uruguay, but they do not investigate whether this improvement in quality was at the expense of quantity. The literature testing the quantity-quality model of fertility (Becker and Lewis, 1973) has found mixed or no evidence of tradeoffs (Black et. al., 2005; Angrist et. al., 2010; Liu, 2014) and family size on quality (Mogstad and Wiswall, 2016). I contribute to this literature by showing that there are no statistically significant changes in the observed health outcomes at birth (birth weight and pregnancy duration).

The remainder of the paper is organized as follows. In Section 2.2, I provide contextual background on the pro-natalist policies in South Korea, primarily focusing on the local pro-natalist cash-transfer programs. Section 2.3 describes the data. I discuss my identification and empirical strategies in Section 2.4 and then present the results in Section 2.5. Section 2.6 concludes with a summary of important findings and recommendations for future research.

2.2 Policy Background

Before the 1960s in South Korea, women on average had 6 children or more. However, the fertility rates of South Korea started to fall in early 60s and have stayed below the 2.1 replacement level since 1983. The decline in the fertility rates did not stop until 2005, when the total fertility rates reached the lowest point at 1.05 children per woman, the lowest in the world according to the World Bank. In 2006, the national pro-natalist movement started with the promulgation of the First Basic Plan for Low Fertility and Aged Society. This was followed by the Second in 2011 and the Third in 2016. As summarized in Lee (2009), the First Basic Plan consisted of five pillars: 1) attenuating the socioeconomic burden of childcare for families with children; 2) expansion of childcare infrastructure; 3) expansion of support for pregnancy and childbirth; 4) increasing compatibility between work and home; and 5) promoting gender-equal family and social culture. Should certain agendas drafted in the Plans be executed, all the districts would be uniformly affected. For instance, one of a few policies that resulted from the Basic Plan is childcare support. Starting from 2013, childcare support has been provided unconditionally to every third child uniformly across South Korea. Moreover, it is important to know that the Basic Plans operated at the national level, and the national government cannot specify policy prescriptions for local governments.[4]

Independent of the national pro-natalist efforts, the local governments have adopted and funded their own cash transfer programs to encourage childbirth since 2001. There are some important empirical facts about the local policy to note. First, one of the remarkable aspects of the local pro-natalist policies in South Korea is that all the programs provide cash transfer differing in generosity by birth parity across localities, and that this is the only policy measure they adopt for their pro-natalist program. In other words, the only benefit local governments provide to their residents after having a newborn baby is cash transfer. Second, cash transfer generosity for third children is representative of pro-natalist policy for each locality. It is the general consensus among the local government officials in charge of pro-natalist cash transfer that cash transfer for the third child is determined first, and then cash transfer generosity for the other parities take the generosity for third child as a benchmark. For the remainder of the paper, I only focus on cash transfers for third children of less than 7,200,000 KRW (or approximately 7,200 USD) when referring to overall cash-transfer generosity, excluding the top 5 percent of the sample with strictly positive cash transfer. Third, since the first implementation of pro-natalist cash transfer in 2001 by 23 districts, all the districts adopted their own pro-natalist cash transfer by 2013. Lastly, the generosity increases over time; the mean cash transfers are 100 USD in 2001, 1,727 USD in 2008 and 2,663 USD in 2015. The local pro-natalist cash-transfer policy adoption rate and the average cash transfer generosity from 2000 to 2015 are plotted in 2.1. In sum, there exist both cross-sectional and within-district over-time variations in cash-transfer generosity.[5]

2.3 Data

In order study the effects of the local pro-natalist cash-transfer programs on birth outcomes, I constructed an annual panel dataset based on multiple administrative sources from 2000 to 2015 for 222 districts that belong to 15 provinces. The panel dataset can be broadly categorized into three components: local pro-natalist cash-transfer information,

[4] As Kim (2013) writes, the Basic Plans launched by the national government "set abstract goals and directions, but did not specify guidelines for local policy formulation."

[5] The Ministry of Family and Welfare reports that the take-up rate of local cash-transfer policy is estimated to be close to 100 percent. Negligible transaction costs arising from simple, convenient application processes may explain high take-up rates (Currie, 2006; Kleven and Kopczuk, 2011). It is required by the law to register every baby within 30 days of birth at a civic center. In every civic center, the sections for birth registry and cash-transfer application are always adjacent to each other. It takes about 10 minutes to complete the required forms to receive cash transfer. To validate, I collected the take up information for 68 districts during the sample period and find that the average take-up rate for 68 districts is 99.98 percent.

number of births and measures of quality, and local characteristics.

2.3.1 Local Cash-Transfer Policy Data

The data on local pro-natalist cash transfer is constructed from three different courses. First, the Ministry of Health and Welfare of South Korea has published the Annual Case Study of Local Government Population Policies since 2008. The case studies provide the information regarding cash-transfer amounts by parity, eligibility, and whether each program is new or continued from the previous year. This is the only official source that allows a direct cross-sectional comparison across districts. However, there is no information prior to 2007 and the case studies are based on the reports voluntarily submitted by local governments. It is often the case that its local pro-natalist cash-transfer details are updated correctly if a district chooses not to share its cash-transfer program with the Ministry. Second, the Enforced Local Laws and Regulations Information System (ELIS), operated by the Ministry of the Interior, is an alternative source for data. This online system provides a vast majority of the laws and regulations enacted by local government entities since 1995 and contains detailed information regarding the local pro-natalist cash-transfer policies. However, at the local level, it is often the case that there exist some discrepancies between the regulation laid out at the time of enactment and their actual implementation. For instance, a local government may not be able to start the program until a few months after its effective date due to lack of administrative or financial resources. Lastly, I exercised the Right to Know under the Official Information Disclosure Act and submitted a formal request for detailed information on its pro-natalist policies to each local government via Open Information System (OIS). In order to ensure the accuracy, I cross-checked the information from the three sources and verified the data with a telephone survey of every local government to resolve any discrepancies. Panel A of Table 2.9 in the Appendix provides the summary statistics of cash transfer by parity.

2.3.2 Quantity and Quality Measures of Birth and Sex Composition Data

I constructed annual data on quantity and quality measures of birth at the district level from the Korean Statistical Information Service (KOSIS) and Microdata Integrated Service (MDIS), both operated by the Bureau of Statistics of South Korea. KOSIS provides annual district-level total fertility rates, the main outcome variable of interest. In order to further decompose number of births by year and by district, I turn to MDIS, which manages all the birth registry records in South Korea. The year, sex, and parity of a birth, district in which records were registered, the ages of parents, birth weight in kilograms, and pregnancy duration in weeks are observed for every record. In order to test whether or not number of births changed due to timing of birth (i.e., delaying birth or bringing it forward), I also extracted the information on the mean ages of parents for all births and by parity. As measures of quantity of birth in additional total fertility rates, I constructed parity-specific and/or age-specific birth rates for each district from 2000 to 2015 based on the birth registry records. To measure "quality", I calculated the mean birth weight, the fraction of low birth weight (less than 2.5 kilograms), the mean pregnancy duration, and the fraction of premature births (fewer than 37 weeks) for all births and by parity. For sex composition at birth, differences in the number of male and female births and female-to-male ratio at birth for all births and by parity were calculated. The summary statistics of the computed quantity and quality measures, and sex composition of birth are presented in Panels B to E of Table 2.9 and Table 2.10.

2.3.3 Local Characteristics Data

In order to control for district characteristics that are potentially correlated with fertility outcomes and local pro-natalist cash transfer, I constructed annual district-level characteristics data from three sources: KOSIS, Finance Integrated System and the National Election Commission. The demographic characteristics include proportion of female population, proportion of adult population, death rate, marriage rate, average female and male ages at first marriage, population density, and net migration per 1,000 people. Government characteristics include per capita budget, party identification, and gender of governing head. Other district characteristics include number of firms, number of laborers, and number of kindergartens. These variables are summarized in Panels F to H of Table 2.11.

2.4 Empirical Strategy

In order to motivate the main empirical strategy to identify the causal effect of local pro-natalist policy on fertility outcomes, I present some preliminary findings based on auxiliary specifications. Exploiting the national pro-natalist policy implementation in 2006 and the variation in the timing of local policy implementation, I estimate the difference in total fertility rates pre-post implementation with district fixed effects. The overall results provide prima facie evidence that the national policy had a positive effect on total fertility rates; the total fertility rates were declining prior to 2006 and reversed this trend thereafter as shown in Figure 2.1. The estimation results are summarized in

Table 2.12. In order to estimate the causal effect of local policy, I introduce year fixed effects in my main empirical model to control for nation-wide policy effects as well as other time-varying aggregate unobservables. Next, based on different implementation timing for each district, I conduct an event study. I semi-parametrically estimate the local policy effects before and after based on the specification exploiting the cross-sectional variation in timing of policy implementation within province as follows:

$$TFR_{d,t} = \gamma_{p,t} + \sum_{\tau=-13}^{13} \gamma_1^{(\tau)} D_{d,t}^{(\tau)} + X'_{d,t-1}\Gamma + \nu_{d,t}, \tag{2.1}$$

where $\left\{D_{d,t}^{(\tau)}\right\}_{\tau=-13}^{13}$ is a set of dummy variables indicating whether or not district d of province p in year t is τ years after its policy implemebtation with $D_{d,t}^{(0)}$ as a leave-out. Figure 2.2 plots the estimated local policy effects before and after implementation. None of the coefficients is statistically significantly different from zero and the standard errors increase when moving away from implementation year, indicated by the 95% confidence intervals based on standard errors clustered at the district level following Bertrand, Duflo, and Mullainathan (2004). All the estimates after implementation are positive and increase in time since implementation. This serves as evidence that the local policy is likely to have a positive effect on total fertility rates. The larger standard errors are likely due to the cross-sectional variations in cash-transfer generosity. To summarize, because each district adopts its policy with varying generosity and the generosity increases after its adoption, it is important to take the variation in cash-transfer generosity into consideration. Therefore, the main empirical model takes generosity as the main explanatory variable to estimate the causal effect of local pro-natalist policy.

The main identification strategy is based on a weighted least squares model with the province-by-year fixed effects, exploiting the cross-sectional variation in cash-transfer generosity within province and year. The female population between the ages of 15 and 49 is used as regression weight, based on the definition of total fertility rates. Naturally, the identifying assumption is that the cash-transfer generosity is exogenous conditional on the observed district-level characteristics by province and year. I now present my main empirical model to estimate the overall local policy effect on total fertility rates at the district level as follows:

$$TFR_{d,t} = \beta_{p,t} + \beta_1 CT_{d,t-1} + X'_{d,t-1}\Gamma + \epsilon_{d,t}, \tag{2.2}$$

where $CT_{d,t-1}$ is a lagged cash transfer (approximately in 1,000 USD) in district d of province p. $X_{d,t-1}$ is a vector of lagged district-level demographics, government, and other characteristics. Standard errors are clustered at the district level, allowing autocorrelation overtime for each district.

A potential threat to identification arises if there are unobserved factors that are correlated with both the cash-transfer generosity and total fertility rates. For instance, a greater level of wealth of a district may allow its local government to provide more generous cash transfer, and a district with a wealthier population may have a higher fertility rate. In this case, the coefficient estimate would be biased upward. Although per-capita government budget and financial independence rate are used as control variables, this problem may persist. Therefore, in order to address this type of threat from omitting unobserved time-invariant district-level characteristics, I exploit the changes in cash transfers within districts over time, instead of the within-province cross-sectional variation in cash transfer, and check whether the estimates of the effects of local policy on total fertility rates are sensitive to this change. In addition, migrations of beneficiaries into and migrations of non-beneficiaries out of districts with more generous cash transfer can overestimate the local policy effect. Because one of the control variables, net migration rate, is a simple difference of inflow and outflow of population, this variable may not reveal much information about how much of inflow or outflow of population is beneficiaries. In order to address this concern, I propose an instrumental variable defined as lagged average cash transfer generosity of adjacent districts in the same province and implement the two-stage least squares method. The relevance condition of the proposed instrument can be tested. The condition for instrument exogeneity is justified by circumventing the simultaneity using the lagged average cash-transfer generosity of contiguous districts in the same province. Furthermore, it is unreasonable to anticipate that the female population of a district responds to the cash transfers of the neighboring districts. Therefore, the exclusion restriction is not likely violated. Given the validity of the instrument, I apply a Durbin-Wu-Hausman specification test of the null hypothesis that the previous WLS estimate of policy effect is not statistically different from the two-stage least squares estimate. Failure to reject the null provides supportive evidence for unbiasedness of the WLS estimate with the province-by-year fixed effect, implying an insignificant degree of biases arising from endogeneity. Lastly, I provide supportive results from a placebo test. For this exercise, I permute the histories of cash transfer across districts within province and estimate placebo effect based on my main empirical model with a “wrong” history of treatment assigned to each district. Repeating the procedure 100,000 times, I nonparametrically estimate the distribution of the placebo policy effect.

After verifying that within-province cross-sectional variations in cash transfers are plausibly exogenous conditional on the observed district-level characteristics, I extend the analysis to study the effects of local policy on birth rates by parity, age of mother, and quality and sex-composition of birth. First, decomposing the total fertility rates by parity and taking advantage of the variations in cash transfer by parity, I separately estimate an econometric model for each parity o for first, second, and third parities in the following form:

$$CBR_{d,t}^{(o)} = \pi_{p,t} + \sum_{o=1}^{3} \pi_1^{(o)} CT_{d,t-1}^{(o)} + X'_{d,t-1}\Gamma + \epsilon_{d,t}^{(o)}, \tag{2.3}$$

where $CBR_{d,t}^{(o)}$ is crude birth rates for parity o in district d of province p in year t and $CT_{d,t-1}^{(o)}$ is cash-transfer generosity for parity o in approximately 1,000 USD. When estimating the policy effect on crude birth rates for parity o, it is reasonable to expect no significant effects from cash transfer generosity for parities less than o. The other more probable spillover is from cash transfers for higher birth parities. For example, if parents with one child are forward looking, they may choose to have a second child so that they would receive the cash transfer for their third child after having their second child. If this were the case, the effect of cash transfer for third children on birth rate of second children would be positive and significant. This explains a potential mechanism through which birth rate of a lower parity is affected by cash transfers for higher parities. Therefore, by including the cash transfers for all parities in the model, I test the extent of spillover effects across parities, and the results show no spillover effects (i.e., the effects of cash transfers are parity-specific).

Second, in order to examine the policy effect on birth rates by age of mother, I regress age-specific birth rates on overall cash-transfer generosity, weighted by the corresponding age-specific female population. Based on the parity specificity of local policy effect, I further decompose the total fertility rates by age of mother and birth parity and study the effects of local policy on age-and-parity-specific birth rates. Third, it may be the case that the cash transfers only have tempo effects rather than quantum effects; local cash transfer may only hasten the timing of pregnancy, as opposed to increasing the number of births. I estimate the effects of cash transfer on female age at first marriage and mother's age structure upon delivery of a child to test whether the cash transfers actually increased the total number of births per woman or simply perturbed the timing without affecting the total number of children a woman would have on average. Fourth, questions regarding possible quantity-quality trade-off are addressed by estimating the effects of cash transfer on birth weight and pregnancy duration as proxies of quality at birth. Lastly, I further access potential impacts of local policy on son preference extant in the context of South Korea using the difference between male and female birth rates and its ratio as dependent variables.

2.5 Results

2.5.1 Local Policy Effects on Quantity of Births

Overall Policy Effect on Total Fertility Rates—Table 2.1 reports the results from estimating the overall effect of local pro-natalist cash transfer on total fertility rates. I start with estimating β_1 in Equation (2) without any fixed effects or the covariates in Column 1 of Panel A and gradually introduce the province fixed effects, year fixed effecst, both fixed effects, and province-by-year fixed effect across columns. The estimate in Column 1 suggests that total fertility rates on average increased by 0.0558 with cash transfer of 1,000 USD. This estimate is biased because the orthogonality of cash transfers and total fertility rates is violated. After controlling for all of time-invariant province-level characteristics with the province fixed effects, the coefficient estimate drops to 0.0323 according to Column 2. However, this estimate is likely still biased upward from positive aggregate shocks that could potentially include the national policy effect discussed earlier. To address this concern, year fixed effects are included and the estimation result indeed shows a moderate drop to 0.0259 in Column 3. Finally, Column 4 reports the local policy effect after completely absorbing all province-level characteristics, both time-varying and time-invariant, by interacting the province and year fixed effects. The estimate is positive at 0.0179, but statistically not significant.

With the province-by-year fixed effects, Equation (2) estimates the local policy effect, effectively comparing fertility rates and cash-transfer generosity across districts within each province for each year. However, the estimates of the coefficient and its standard error for the local policy effect suffer from omitting district-level characteristics that are potentially correlated with both cash-transfer generosity and total fertility rates. For instance, a district with a high population density may have less generous cash transfer, but have a higher fertility rate. Moreover, a district with a governing head who belongs to a conservative party is associated with less generous cash transfer. If people who vote for conservative party members are likely to have preference for greater number of children, then excluding the information about party identification of governing heads results in underestimation of the local policy effect. In Panel B, I report the estimates of the local policy effects with the full set of province-by-year fixed effects, gradually adding observed district-level demographics, government, and other characteristics across columns. Column 2 presents the policy effect after introducing a set of demographic characteristics. The estimate is significant at the 5-percent significance level and indicates a strong positive local policy effect. Including a set of government characteristics in Column 3 and that of other local characteristics in Column 4, the estimated policy effects do not change in a meaningful way and are still statistically significant. In conclusion, according to the coefficient estimate from the fully saturated model in Column 4, a cash transfer of 1,000 USD increased total fertility rates by 0.0216 children per woman or 1.8 percent on average.[6]

I test the robustness of the causal policy effect identified based on the cross-sectional variation in the generosity of cash transfer by estimating the local policy effect based on two alternative models. The estimate of local policy effect on total fertility rate is potentially biased if there are unobserved time-varying district-level characteristics that are correlated with the generosity of cash transfer and total fertility rates. First, I introduce district and year fixed effects to the fully saturated model instead of province-by-year fixed effects. The identifying assumption for this approach is that changes in total fertility rate and in cash-transfer generosity are uncorrelated with changes in district-level unobservables conditional on a set of observed characteristics. Second, I employ an instrumental variables approach using the lagged average cash-transfer generosity of neighboring districts in the same province as an instrument given the plausible validity of the instrument. The estimates of the policy effect based on two alternative models are summarized in Table 2.2. I report the previous estimate based on the preferred cross-sectional variation in Column 1 of Table 2.2. In Column 2, the local policy effect is identified off of the changes in cash-transfer generosity over time within district. The coefficient is significant and suggests that total fertility rate increases by 0.0221 children per woman with a cash transfer in the amount of 1,000 USD.

The last three columns of Table 2.2 report the estimates based on the instrumental variable approach. The reduced form estimate in Column 3 suggests that the lagged average cash-transfer generosity of neighboring districts increased total fertility rate by 0.0118 children per woman. According to Column 4, the first stage coefficient is positive and statistically significant, confirming the relevance condition for the instrument. This result implies an increase in cash transfer by 412 USD in response to a 1,000 USD increase in the lagged average cash transfer of neighboring districts. In Column 5, the estimate based on the two-stage least squares approach is 0.0288, statistically significant. Applying a Durbin-Wu-Hausman specification test, I find an F-statistic equal to 0.76 and cannot reject the null hypothesis at 1-, 5-, and 10-percent significance levels (p-value = 0.3841), suggesting that there is no statistical difference between the estimates from weighted least squares and two-stage least squares methods. Lastly, Figure 2.4 plots the

[6] Figure 2.3 in the Appendix presents a set of scatter plots in which each circle corresponds to a residual cash-transfer generosity on the horizontal axis and residual total fertility rate on the vertical axis from unweighted and weighted regressions by female population between the ages of 15 and 49. Size of hollow circles reflects the female population between the ages of 15 and 49. The residuals are estimated based on the preferred model in Equation (2) with the full set of observed district characteristics.

distribution of placebo policy effect, nonparametrically estimated from 100,000 iterations of permuting the treatment histories among districts within provinces and running the preferred empirical model. I reject the null hypothesis that the estimated policy effect in Column 1 of Table2.2 and the placebo effect are the same at the 1-percent significance level (p-value = 0.0001). Cross-examining the estimates of local policy effect on total fertility rates from different models, I conclude that the estimate of causal local policy effect from the preferred model using the within-province cross-sectional variation in cash transfer (Column 1) is robust to different identification assumptions and estimation methods, therefore plausibly unbiased.[7]

Birth Rates by Parity and Age of Mother— Table 2.3 presents regressions estimating the effect of cash transfer on parity-specific birth rates. Column 1 estimates the effect of cash transfer for first children on birth rates for first children based on Equation (3) by assuming no spillover effects across parities (i.e. $\pi_1^{(2)} = \pi_1^{(3)} = 0$) . The coefficient estimate is positive and statistically significant and suggests that birth rates for first children increased by 2.771 first children per 1,000 women or 15.56 percent with a cash transfer of 1,000 USD for first children. Should parents be forward looking and base their decision to have a first child on the benefits offered for higher birth orders, birth rates for first children would be affected by cash transfer for high birth orders. If this is indeed the case, cash transfers for second and third children would have positive effects on birth rates for first children. To test whether this is the case, I estimate Equation (3) without the no-spillover restrictions in Column 2. While the coefficient estimate for cash transfer for first children does not change in a meaningful way and stays statistically significant, the estimated effects of cash transfers for higher-order births are statistically not different from zero. This implies that the cash-transfer effect is parity-specific for first children and that the cash transfers did not inframarinally affect fertility. Moreover, this result rules out a potential mechanism that birth rates for first children increased due to increased births of unwanted babies as an unintended consequence of cash incentives brought by the local pro-natalist programs.

The results for crude birth rates for second children are very much line with the previous results. The effect of cash transfer for second children is consistently estimated in Column 3 and Column 4 of Table 2.3. According to the estimation result in Column 3, a cash transfer of 1,000 USD increased birth rate for second children by 1.353 births per 1,000 women or 9.81 percent. Column 4 indicates that birth rates for second children were

[7] In the Appendix, I provide additional evidence that there are no other district characteristics changing along with birth rates and cash transfers. Figure 2.6 in the Appendix presents a set of figures plotting the nonparametric estimates of selected district characteristics before and after local policy implementation. Table 2.13 reports the estimated effects of cash-transfer generosity on selected district characteristics.

unaffected by the cash transfer for third children. Furthermore, the coefficient for cash transfer for first children is also insignificant. This serves as a falsification test because it is obvious to expect that cash transfers for first children does not affect birth rates of second children. Lastly, the estimation results for birth rates for third children are positive, statistically significant, and consistently estimated in Column 5 and Column 6; it suggests that a cash transfer of 1,000 USD increased crude birth rates for third children by 0.916 births per 1,000 women or 3.06 percent. Once again, the results in Column 6 show that cash transfer for first and second children did not have an impact on birth rates for third children as the coefficient estimates for cash transfers for first and second children are not significantly different from zero. The results in Table 2.3 altogether imply that the cash-transfer effects were parity-specific and that the magnitude of policy effect decreases with birth parity.

Overall local policy effects on birth rates by age of mother are reported in Table 2.4. Column 1 takes crude birth rates of the female population between the ages of 15 and 19 as the dependent variable. The estimate of policy effect is negative, but close to zero and statistically insignificant. This indicates that the cash incentive provided by local governments did not affect adolescent birth rates. From Column 2 to Column 7, the estimated overall policy effect on birth rates of the female population for each age group is positive, increasing with age of the female population younger than 35 and decreasing thereafter. However, the estimates are statistically significant at 10 percent except the one in Column 2. The standard error for the estimated policy effect in Column 2 is large almost certainly due to the fact that there are generally much less mothers between the ages of 20 to 24 with two children whose decision to have a third child would be affected by cash transfer for third child. Although positive, the results in Column 6 and Column 7 suggest that the policy effects on birth rates of female population between the ages of 40 to 49 are negligible. This result comes as no surprise because the female population in these age groups are likely to have completed their fertility decisions.

In light of the previous finding that cash transfers have parity-specific effects, Table 2.5 completely characterizes the local policy effects by parity and the female population between the ages of 20 and 39, who are more active in their fertility decisions. In Panel A, age-specific birth rates for first children are regressed on cash-transfer generosity for first children. The estimates are positive and statistically significant across columns, implying that the cash transfer is most effective for females between the ages of 25 and 29. Panel B reports the results for second children. While all the estimates are positive and statistically significant across age groups, cash transfers for second children had greater effects for the female population between the ages of 25 to 34. Finally, the results assessing the effects on birth rates for third children indicate that the cash-transfer policies were effective and relevant for the female population between the ages of 30 and 39. Consistent with the previous results on parity-specific birth rates, the estimated policy effects in Table 2.5 decrease with parity for each age group.

Female Age at First Marriage and Age of Mother at Birth— Table 2.6 provides evidence that the positive effects of cash transfer on quantity of births are not driven by timing and spacing of births. In Panel A, I assess the effects of cash transfer by parity on average female age at first marriage. Although marriage is not required for receiving cash transfer, average female age at first marriage is likely to decrease in cash-transfer generosity if potential mothers were simply to actualize their fertility plans early, thus more likely to get married at an earlier age. In Column 1, the estimated effect of cash transfer for first children on average female age at first marriage is positive and significant at the 10-percent level. The result implies that a cash transfer of 1,000 USD lowered the average female age by 19 days. The estimates in Column 2 and Column 3 are closer to zero and statistically not different from zero, suggesting that the cash transfers for second and third children did not have impacts on average female age at first marriage. Column 4 includes cash transfers for first, second, and third children as explanatory variables and finds that only cash transfers for first children have positive and significant effect as before. The results presented in Panel A of Table 2.6 indicates that local pro-natalist cash transfer did not result in early marriage for the female population, which could have also brought fertility forward in time. Panel B investigates whether or not the cash-transfer program had any impacts on the average age of mothers. The majority of the estimates are negative; mothers were on average younger by 1 to 28 days due to a cash transfer of 1,000 USD. However, none of the coefficients is significantly different from zero across columns. Panel B provides strong evidence against a mere tempo effect of local cash transfers. The results reported in Table 2.6 suggest that the increase in number of births was not driven by merely adjusting the timing of birth forward and reducing the duration of time between births.

2.5.2 Local Policy Effects on Health Outcomes at Birth and Son Preference

Birth Weights and Pregnancy Duration—I assess the local policy effects of pro-natalist cash transfer on quality measures of birth and report the results in Table 2.7. First, the estimate in Column 1 of Panel A indicates no evidence on overall changes in average birth weights in kilograms due to cash transfer. When investigating the effects by birth parity in Column 2, Column 3, and Column 4, I only find a negative yet significant effect for second children. However, while significant, the estimated effect for second children suggests a decrease in birth weight only by 5.51 grams or by 0.012 percent. In addition, all of the estimates presented in Panel B indicate no change in number of incidences of birth with low birth weight, defined by less than 2.5 kilograms. Another quality measure of birth observed is pregnancy duration in weeks. I report the estimated policy effects on average pregnancy duration (in Panel C) and fraction of premature births, defined by 37 weeks or earlier (in Panel D). None of the results is significantly different from zero, suggesting that pregnancy duration was not impacted by cash transfer. The estimation results presented in Table 2.7 should be interpreted with caution. I acknowledge that the results do not speak to improvement in quality at birth. This is because it is not certain what the optimal levels of birth weight and pregnancy duration are. Unless they are known, it is not possible to discern whether or not a positive or negative coefficient means an increase or a decrease in quality of birth. However, provided that the coefficients are close to, and statistically indistinguishable from zero, I claim that there were no meaningful changes in birth weight and pregnancy duration of births and thus conclude no quality improvement or decline due to cash transfer.

Sex Composition—As Edlund and Lee (2013) find, preference for sons in South Korea is still prevalent in the 21st century, particularly for higher-order births. Given this general proclivity for sons, I attempt to answer whether or not this tendency was intensified by the local pro-natalist cash-transfer programs. The effects of cash transfer on sex composition at birth are evaluated in Table 2.8. Column 1 of Panel A examines the overall policy effect on the difference between male and female crude birth rates. The estimate is negative, which implies that the difference in birth rates between male and female decreases, yet is statistically not different from zero. The estimated effects for first and second children in Column 2 and Column 3 are not statistically different from zero. However, the policy effect is positive and significant for the gap between the male and female birth rates for third children. This serves as additional evidence for son preference already extant for third children, but not for changes in son preference. In Panel B, I estimate the effects on female to male ratio at birth. The estimated coefficients across columns, each indicating effects for overall, first, second, and third births, are statistically not different from zero. To sum up the results on son preference, there indeed existed son preference for third children, but I find no suggestive evidence that cash transfer exacerbated sex composition due to son preference.

2.6 Discussion and Conclusion

In this paper, I study the effects of pro-natalist cash-transfer policy on quantity and quality outcomes of birth. To do so, I collected detailed policy information on local cash-transfer programs for all the districts in South Korea while verifying its accuracy from four different sources. This paper improves the identification of policy effects on fertility rates and other characteristics of births by exploiting the episodes of pro-natalist cash-transfer policy adoption and changes in cash-transfer generosity for each parity at the local level. I corroborated the consistency of the estimated policy effect on total fertility rates by re-estimating the policy parameter based on two alternative models and implementing a placebo test. In Appendix, I provide a set of figures (Figure 2.6) and a table (Table 2.13) that serve as additional robustness checks to illustrate that no relevant variables systematically changed along with adoption and generosity of cash transfer and fertility rates.

I find overall positive causal effects of cash transfer on quantity measures of birth and no meaningful effects on quality and sex composition of birth. Furthermore, cash transfer for each parity only affected the birth rates of the corresponding parity. For instance, cash transfer for second children did not increase birth rates for first and third children. While birth rates of adolescence and women older than 40 were not affected, cash transfer had positive effects on birth rates of the female population between the ages of 20 and 39, who are more active in their fertility decisions. Cash transfer had no meaningful effects on average female age at first marriage and average ages of mothers. This suggests that the estimated policy effects were mostly driven by the increase in completed fertility, rather than a temporary increase in births due to changes in timing of childbearing. Lastly, I find no evidence that local cash-transfer policies changed quality measures of birth and tendency to favor sons over daughters for higher-order births.

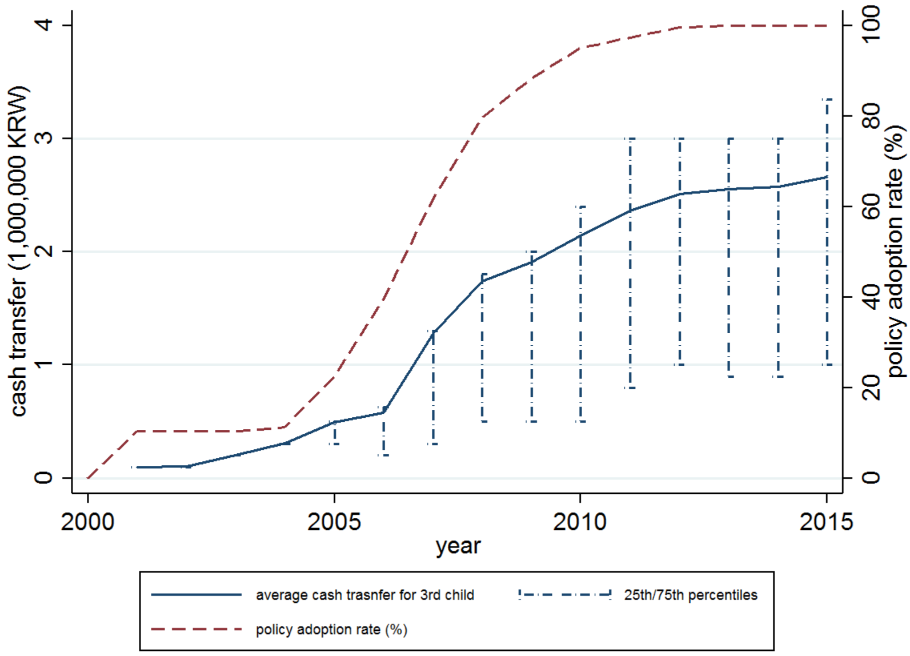
A cash transfer of 1,000 USD increased total fertility rates by 0.0216 children per woman, ceteris paribus. This implies that the benefit elasticity of fertility evaluated at the mean of strictly positive cash transfers (2,042 USD) and the mean total fertility rate (1.218 children per woman) is about 0.036. In other words, a 1 percent increase in cash transfer raised the total fertility rate by 0.036 percent. This implies elasticity is rather a conservative estimate due to potential downward bias from strict eligibility and imperfect take-ups. However, this potential bias is unlikely because, as discussed earlier, the eligibility condition is satisfied for most cases, as the only requirement is residency and the observed take-up rate is 100 percent. On the one hand, the implied elasticity from this paper is smaller than the estimates of benefit elasticity from the similar papers, which computed the elasticity of fertility from various forms of financial incentives to range from 0.05 to 0.248, summarized in Cohen et. al. (2013). On the other hand, when evaluated at the mean annual disposable household income in South Korea (19,372 USD in 2014 according to OECD's Better Life Index), the elasticity is 0.34, consistent with the related estimates of 0.3-0.4 in Black et. al. (2008) and theoretical prediction of positive income elasticity for higher-income countries in Becker and Thomas (1976).

Although there have been some efforts to assess the effectiveness of pro-natalist policies apart from cash transfers, the results from the studies that examine the effects of parental leave and tax benefits on fertility are often mixed (Lalive and Zweimüller, 2009; Andersson and Duvander, 2006; Whittington et. al., 1990; Crump et. al., 2011). The overall results of the paper provide convincing evidence that cash transfer is an effective form of policy measure to increase completed fertility and number of children ever born per woman without changing the quality and sex composition at birth. There are some important issues unaddressed in this paper, which stipulate separate future research. First, as it was the case in the context of the pro-natalist policies in South Korea, it is often the case that the interests of local governments and their federal government align, each implementing own policies. It is uncertain whether the policies administrated at different tiers of government reinforce or undermine each other. Second, this paper skirted an important normative question. What is the optimum level of population? Third, a greater number of births means there are more mothers and fathers taking time off from work if not exiting the labor market. Lastly, although there were no apparent short-run effects of cash transfer on quality measures at birth, the cash transfer may change the level of parental investment in children's human capital, which, in the long run, would accumulate to educational and labor-market outcomes.

2.7 Figures and Tables

Figure 2.1: Local Pro-natalist Policy and Total Fertility Rate over Time

(a) Cash Transfer Generosity and Prevalence



(b) Cash Transfer Prevalence vs. TFR

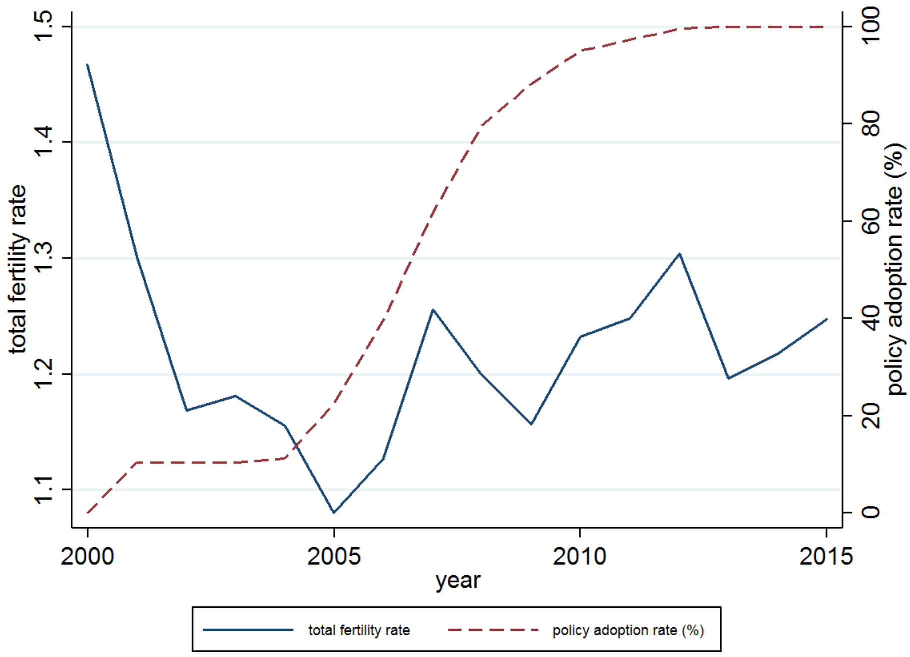


Figure 2.2: Total Fertility Rates Before and After Local Policy Implementation

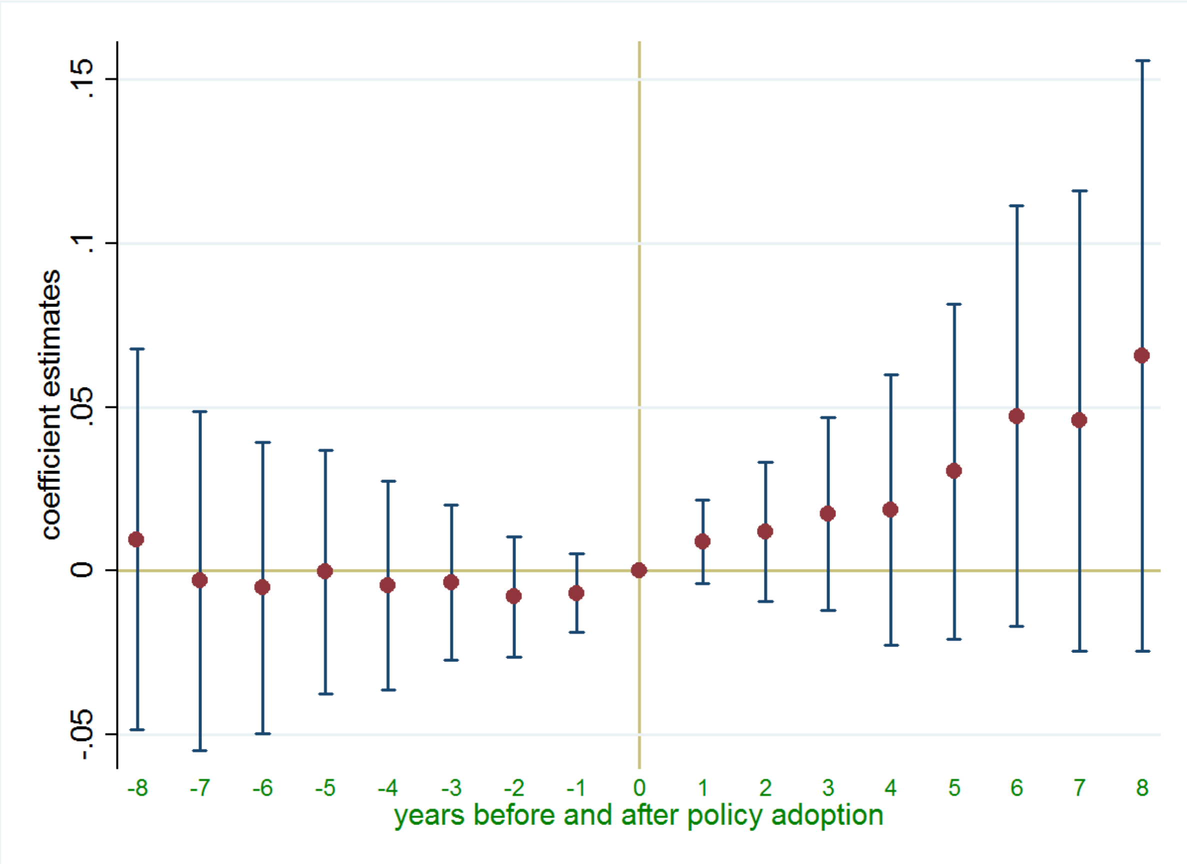


Table 2.1: The Effect of Local Cash Transfer on Total Fertility Rates

Dependent Variable: Total Fertility Rates ($TFR_{d,p,t}$)	(1)	(2)	(3)	(4)
A. Fixed Effects				
Cash Transfer ($CT_{d,p,t-1}$)	0.0558***	0.0323***	0.0259***	0.0179
	(0.00923)	(0.00697)	(0.00990)	(0.0137)
Observations	3,256	3,256	3,256	3,256
R^2	0.057	0.503	0.572	0.598
Province FE	N	Y	Y	N
Year FE	N	N	Y	N
Province-by-Year FE	N	N	N	Y
B. Control Variables				
Cash Transfer ($CT_{d,p,t-1}$)	0.0179	0.0255**	0.0246**	0.0216**
	(0.0137)	(0.0101)	(0.00993)	(0.00958)
Observations	3,256	3,256	3,253	3,246
R^2	0.598	0.840	0.843	0.850
Mean Dependent Variable	1.218	1.218	1.218	1.218
Province-by-Year FE	Y	Y	Y	Y
Control for:				
Demographic Characteristics	N	Y	Y	Y
Governmental Characteristics	N	N	Y	Y
Other Characteristics	N	N	N	Y

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by female population between the ages of 15 and 49; (ii) Demographic characteristics include proportion of female population, proportion of adult population, death rate, marriage rate, ages at first marriage for men and women, and log of population density and net migration rate; (iii) Government characteristics include per-capita budget, financial independence rate, party identification and gender of governing head, and land trade rate; (iv) Other characteristics include number of firms and number of laborers per 1,000 people and number of kindergartens per 1,000 children

Table 2.2: Local Policy Effects on Total Fertility Rates Based on Alternative Models

	(1)	(2)	(3)	(4)	(5)
	WLS	WLS	Reduced	First	2SLS
Dependent variables:	$TFR_{d,p,t}$	$TFR_{d,p,t}$	$TFR_{d,p,t}$	$CT_{d,p,t-1}$	$TFR_{d,p,t}$
Cash Transfer ($CT_{d,p,t-1}$)	0.0216**	0.0221***			0.0288***
	(0.00958)	(0.00459)			(0.00846)
Avg. CT of Adjacent Districts			0.0118***	0.412***	
			(0.00426)	(0.0703)	
Observations	3,246	3,246	3,246	3,246	3,246
R^2	0.850	0.931	0.929	0.717	0.930
Mean Dependent Variable	1.218	1.218	1.218	1.218	1.218
Province-by-Year FE	Y	N	N	N	N
Year FE	N	Y	Y	Y	Y
District FE	N	Y	Y	Y	Y

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by the total number of births of the corresponding parity; (ii) Across columns, the same set of control variables are used as the fully specified model in Column 4 of Panel B of Table 2.1

Table 2.3: The Local Cash Transfer Effects by Parity

Dependent variable:	(1)	(2)	(3)	(4)	(5)	(6)
Fertility rates for parity	1st	1st	2nd	2nd	3rd	3rd
Cash Transfer for 1st child	2.771***	2.947***		0.409		0.0334
	(0.496)	(0.511)		(0.553)		(0.123)
Cash Transfer for 2nd child		-0.167	1.353***	1.303***		0.0160
		(0.212)	(0.272)	(0.379)		(0.0738)
Cash Transfer for 3rd child		0.00813		-0.0273	0.0916***	0.0830**
		(0.0641)		(0.0669)	(0.0259)	(0.0332)
Observations	3,279	3,279	3,234	3,234	3,246	3,246
R^2	0.954	0.954	0.932	0.932	0.920	0.920
Mean Dependent Variable	17.8009	17.8009	13.7967	13.7967	2.9897	2.9897

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by female population between the ages of 15 and 49; (ii) Across columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1; (iii) I exclude the observations with cash transfers for first, second, and third children above the 95th percentile.

Table 2.4: The Local Cash Transfer Effects by Parity and Age of Mother

	Age Group of Mother						
Dependent variable:	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Fertility rates for age group:	15-19	20-24	25-29	30-34	35-39	40-45	45-49
Cash Transfer ($CT_{d,p,t-1}$)	-0.00665	0.335	1.657*	1.802*	0.345*	0.0113	0.000526
	(0.0346)	(0.234)	(0.846)	(0.974)	(0.204)	(0.0366)	(0.00399)
Observations	3,246	3,246	3,246	3,246	3,246	3,246	3,246
R^2	0.700	0.905	0.926	0.887	0.922	0.807	0.229
Mean Dependent Variable	1.4489	19.5648	92.6862	95.5670	26.8431	3.3446	0.1618

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by female population of the corresponding age groups; (ii) Across columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1.

Table 2.5: The Local Policy Effects by Parity and Age of Mother

	Age Group of Mother			
	(1)	(2)	(3)	(4)
	20-24	25-29	30-34	35-39
Dependent Variable: Age-Specific Birth Rates of First Child				
Cash Transfer for 1st Child	1.731**	7.835***	4.584***	0.972***
	(0.769)	(2.584)	(1.759)	(0.441)
observations	3,279	3,279	3,279	3,279
R^2	0.897	0.914	0.905	0.913
Mean Dependent Variable	14.6809	58.7699	40.1044	8.2664
Dependent Variable: Age-Specific Birth Rates of Second Child				
Cash Transfer for 2nd Child	0.675***	3.368***	3.544***	0.701**
	(0.178)	(1.023)	(1.193)	(0.341)
observations	3,234	3,234	3,234	3,234
R^2	0.849	0.917	0.852	0.907
Mean Dependent Variable	4.4935	30.3459	45.1602	12.0937
Dependent Variable: Age-Specific Birth Rates of Third Child				
Cash Transfer for 3rd Child	0.0171	0.0611	0.323***	0.164**
	(0.0165)	(0.0535)	(0.0986)	(0.0526)
observations	3,246	3,246	3,246	3,246
R^2	0.559	0.857	0.877	0.806
Mean Dependent Variable	0.3698	3.3539	9.5182	5.6079

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by female population of the corresponding age groups; (ii) Across columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1; (iii) I exclude the observations with cash transfers for first, second, third children above the 95th percentile.

Table 2.6: Local Policy Cash Transfer Effects on Timings of Marriage and Birth

	(1)	(2)	(3)	(4)
A. Dependent variable: Female Age at First marriage				
Cash Transfer for 1st Child	0.0523*			0.0393*
	(0.0281)			(0.0222)
Cash Transfer for 2nd Child		-0.0117		-0.0330
		(0.0167)		(0.0204)
Cash Transfer for 3rd Child			0.00418	0.0108
			(0.00502)	(0.00739)
Observations	3,279	3,234	3,246	3,246
R^2	0.980	0.980	0.980	0.980
Mean Dependent Variable	28.4171	28.4171	28.4171	28.4171
B. Dependent Variable:				
Age of Mother at Birth for Parity:	**Overall**	**1st**	**2nd**	**3rd**
Cash Transfer for 1st Child	-0.00440	0.0765	-0.0101	-0.0218
	(0.0448)	(0.0623)	(0.0637)	(0.0760)
Cash Transfer for 2nd Child	-0.0207	-0.0300	-0.0407	-0.0230
	(0.0374)	(0.0365)	(0.0401)	(0.0589)
Cash Transfer for 3rd Child	-0.00731	-0.00365	-0.0106	-0.0189
	(0.0134)	(0.00936)	(0.0101)	(0.0160)
Observations	3,246	3,279	3,234	3,246
R^2	0.972	0.972	0.970	0.861
Mean Dependent Variable	30.1491	29.0136	30.8710	33.0159

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by total number of first marriage for women for Panel A and by the total number of births of the corresponding parity for Panel B; (ii) Across columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1; (iii) I exclude the observations with cash transfers for first, second, and third child above the 95th percentile.

Table 2.7: Local Policy Effects on Birth Weight and Pregnancy Duration

	(1) Overall	(2) 1st	(3) 2nd	(4) 3rd	(5) Overall	(6) 1st	(7) 2nd	(8) 3rd
	A. Average Birth Weight				**B. Fraction of Births with Low Birth Weight**			
Cash Transfer for 1st Child		-0.00175				-0.000428		
		(0.00441)				(0.00112)		
Cash Transfer for 2nd Child			-0.00551*				0.000515	
			(0.00284)				(0.000852)	
Cash Transfer for 3rd Child	-0.000536			0.000754	0.000290			0.000697
	(0.00104)			(0.00144)	(0.000231)			(0.000576)
R^2	0.702	0.630	0.592	0.425	0.420	0.274	0.364	0.183
Mean Dependent Variable	3.2366	3.2285	3.2412	3.2285	0.0519	0.0514	0.0506	0.0577
	C. Average Pregnancy Duration				**D. Fraction of Premature Births**			
Cash Transfer for 1st Child		0.0102				-0.000559		
		(0.0305)				(0.00161)		
Cash Transfer for 2nd Child			0.0115				-0.000854	
			(0.0187)				(0.00129)	
Cash Transfer for 3rd Child	0.00450			0.00182	0.000539			0.000767
	(0.00603)			(0.00890)	(0.000453)			(0.000884)
R^2	0.787	0.733	0.758	0.627	0.651	0.506	0.559	0.317
Mean Dependent Variable	38.8479	39.0716	38.6324	38.5589	0.0526	0.0475	0.0551	0.0671
Observations	3,246	3,279	3,234	3,246	3,246	3,279	3,234	3,246

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by the total number of births of the corresponding parity; (ii) Across columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1; (iii) I exclude the observations with cash transfers for first, second, and third children above the 95th percentile.

Table 2.8: Local Policy Effects Son Preference

	(1) Overall	(2) 1st	(3) 2nd	(4) 3rd
A. Difference of Gender-Specific Fertility Rates				
Cash Transfer for 1st Child		0.0679		
		(0.0619)		
Cash Transfer for 2nd Child			0.0199	
			(0.0286)	
Cash Transfer for 3rd Child	-0.00237			0.0114**
	(0.00613)			(0.00577)
R^2	0.417	0.278	0.324	0.530
Mean Dependent Variable	0.3221	0.3491	0.3119	0.2414
B. Female to Male Ratio at Birth				
Cash Transfer for 1st Child		-0.00736		
		(0.0123)		
Cash Transfer for 2nd Child			0.00380	
			(0.00617)	
Cash Transfer for 3rd Child	0.00102			-0.00349
	(0.00134)			(0.00473)
R^2	0.175	0.087	0.084	0.375
Mean Dependent Variable	0.9348	0.9502	0.9464	0.8397
Observations	3,246	3,279	3,234	3,246

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by the total number of births of the corresponding parity; (ii) Across columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1; (iii) I exclude the observations with cash transfers for first, second, third children above the 95th percentile.

2.8 Appendix A: Additional Figures and Tables

Table 2.9: Summary Statistics (A/B)

Variables	(1) N	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max
A. Cash Transfer Generosity					
Cash Transfer for 1st Child	3,552	0.0409	0.216	0	5.100
Cash Transfer for 2nd Child	3,552	0.170	0.437	0	7.540
Cash Transfer for 3rd Child	3,552	0.609	1.177	0	18.80
Cash Transfer for 4th Child	3,552	0.804	1.585	0	18.80
Cash Transfer for 5th Child	3,552	0.917	1.857	0	24
B. Quantity Measures at Birth					
Total Fertility Rate	3,552	1.218	0.220	0.696	2.470
Birth Rate of 1st Child	3,552	17.80	5.982	2.972	149.1
Birth Rate of 2nd Child	3,552	13.80	5.224	2.566	106.2
Birth Rate of 3rd Child	3,552	2.990	1.492	0.592	26.79
Birth Rate for Female Age 15-19	3,552	1.400	1.159	0	27.55
Birth Rate for Female Age 20-24	3,552	19.45	14.27	0	410.8
Birth Rate for Female Age 25-29	3,552	92.91	47.80	18.51	853.7
Birth Rate for Female Age 30-34	3,552	96.35	36.23	11.68	1,062
Birth Rate for Female Age 35-39	3,552	26.95	13.05	2.732	352.7
Birth Rate for Female Age 40-44	3,552	3.312	1.613	0	32.17
Birth Rate for Female Age 45-	3,552	0.166	0.168	0	3.924
Birth Rate of 1st Birth for Female Age 20-24	3,552	14.63	10.18	0	261.4
Birth Rate of 2nd Birth for Female Age 20-24	3,552	4.443	3.999	0	138.9
Birth Rate of 3rd Birth for Female Age 20-24	3,552	0.361	0.491	0	13.88
Birth Rate of 1st Birth for Female Age 25-29	3,552	58.86	26.40	10.46	529.6
Birth Rate of 2nd Birth for Female Age 25-29	3,552	30.38	20.72	3.145	359.1
Birth Rate of 3rd Birth for Female Age 25-29	3,552	3.449	3.132	0	50.89
Birth Rate of 1st Birth for Female Age 30-34	3,552	40.48	18.93	2.910	500.3
Birth Rate of 2nd Birth for Female Age 30-34	3,552	45.38	16.74	6.018	474.4
Birth Rate of 3rd Birth for Female Age 30-34	3,552	9.668	5.559	0	106.8
Birth Rate of 1st Birth for Female Age 35-39	3,552	8.326	4.900	0	117.5
Birth Rate of 2nd Birth for Female Age 35-39	3,552	12.11	6.364	0	153.5
Birth Rate of 3rd Birth for Female Age 35-39	3,552	5.621	2.564	0	72.42

Table 2.10: Summary Statistics (C/D/E)

Variables	(1) N	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max
C. Quality Measures at Birth					
Birth Weight	3,552	3.235	0.0292	3.070	3.404
Birth Weight for 1st Child	3,552	3.228	0.0297	3.036	3.426
Birth Weight for 2nd Child	3,552	3.240	0.0325	3.021	3.442
Birth Weight for 3rd Child	3,552	3.261	0.0520	2.573	3.767
Fraction of Low Birth Weight	3,552	0.0522	0.00765	0	0.121
Fraction of Low Birth Weight for 1st Child	3,552	0.0516	0.00865	0	0.146
Fraction of Low Birth Weight for 2nd Child	3,552	0.0512	0.0104	0	0.141
Fraction of Low Birth Weight for 3rd Child	3,552	0.0589	0.0191	0	0.333
Pregnancy Duration	3,552	38.84	0.231	37.26	39.71
Pregnancy Duration for 1st Child	3,552	39.07	0.208	37.50	39.87
Pregnancy Duration for 2nd Child	3,552	38.61	0.268	35.67	39.69
Pregnancy Duration for 3rd Child	3,552	38.54	0.338	35.17	39.77
Fraction of Premature Birth	3,552	0.0531	0.0113	0	0.140
Fraction of Premature Birth for 1st Child	3,552	0.0477	0.0106	0	0.145
Fraction of Premature Birth for 2nd Child	3,552	0.0560	0.0142	0	0.227
Fraction of Premature Birth for 3rd Child	3,552	0.0686	0.0236	0	0.333
D. Gender Composition at Birth					
Difference of Male and Female Births	3,552	0.253	0.332	-2.232	2.510
Difference of 1st Male and Female Births	3,552	0.318	0.355	-2.564	3.531
Difference of 2nd Male and Female Births	3,552	0.270	0.329	-3.773	3.319
Difference of 3rd Male and Female Births	3,552	0.200	0.244	-1.420	3.528
Female to Male Ratio	3,552	0.935	0.0465	0.618	1.541
Female to Male Ratio for 1st Birth	3,552	0.950	0.0648	0.519	1.889
Female to Male Ratio for 2nd Birth	3,552	0.947	0.0712	0.375	1.944
Female to Male Ratio for 3rd Birth	3,552	0.839	0.175	0	4.500
E. Age of Mother at Birth					
Age of Mother	3,552	30.24	1.121	27.16	32.86
Age of Mother for 1st Birth	3,552	29.04	1.252	24.85	32.08
Age of Mother for 2nd Birth	3,552	31.03	1.218	27.32	34.06
Age of Mother for 3rd Birth	3,552	33.23	0.954	29.05	38

Table 2.11: Summary Statistics (F/G/H)

Variables	(1) N	(2) Mean	(3) Std. Dev.	(4) Min	(5) Max
F. Demographic Characteristics					
% Female Population	3,552	0.414	0.0214	0.337	0.487
% Adult Population	3,552	0.497	0.0459	0.260	0.580
Death Rate	3,330	4.888	2.226	1.472	47.17
Marriage Rate	3,552	0.0128	0.00378	0.00247	0.103
Female Age at First Marriage	3,552	28.39	1.299	24.09	31.62
Male Age at First Marriage	3,552	30.96	1.078	28.08	35.26
Log(Population Density)	3,552	8.044	1.737	2.968	10.30
Net Migration per 1000 People	3,552	0.0266	14.14	-83.47	147.6
G. Local Government Characteristics					
Financial Independence Rate	3,330	40.60	17.55	6.400	95.30
Per Capita Budget	3,327	1.462	1.525	0.239	21.87
Conservative Party	3,330	0.626	0.484	0	1
Female Government Head	3,330	0.0171	0.130	0	1
Land Trader per 1000 People	3,323	41.83	88.63	0.640	1,372
H. Other Characteristics					
# Firms per 1000 People	3,330	66.46	30.98	16.20	528.5
# Laborers per 1000 People	3,330	330.0	207.2	64.82	2,970
# Kindergartens per 1000 Children	3,330	0.673	0.452	0.179	5.184

Table 2.12: Diff-In-Diff Estimates of Pro-natalist Policy Effect on TFR

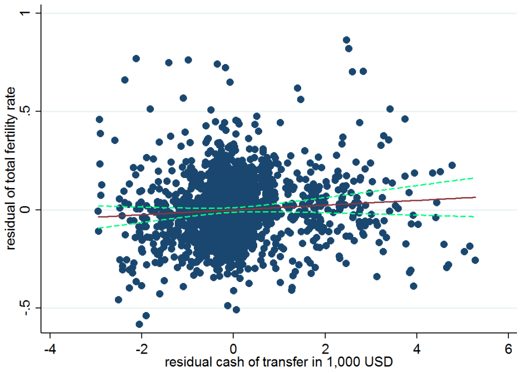
Dependent variable: Total Fertility Rate ($TFR_{c,p,t}$)	(1) National	(2) Local	(3) Interaction
Post National Policy Implementation (P_t)	0.174***		0.170***
	(0.00689)		(0.00723)
Post Local Policy Implementation ($A_{d,p,t}$)		-0.00823	-0.0365*
		(0.00643)	(0.0204)
$P_t \times A_{d,p,t}$			0.0315
			(0.0207)
Observations	3,320	3,320	3,320
R-squared	0.907	0.893	0.907
District FE	Yes	Yes	Yes
Control for District Characteristics	Yes	Yes	Yes
Control for Province and National Trends	Yes	Yes	Yes

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1. Weighted by female population between the ages of 15-49; (ii) District characteristics include demographic, government, and other characteristics listed in Table 2.11; (iii) Province and national trend control variables include female and male unemployment and labor force participation rates, gross regional domestic product (province-level), and number of new houses (province-level).

Figure 2.3: Residual Analysis: Weighted vs. Unweighted

A. Unweighted

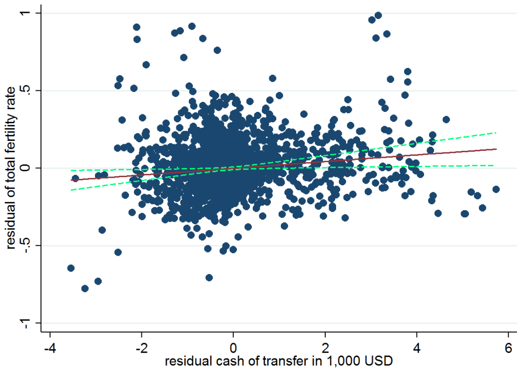
(a)



(b)

B. Weighted by Female Population between the Ages of 15 and 49

(c)



(d)

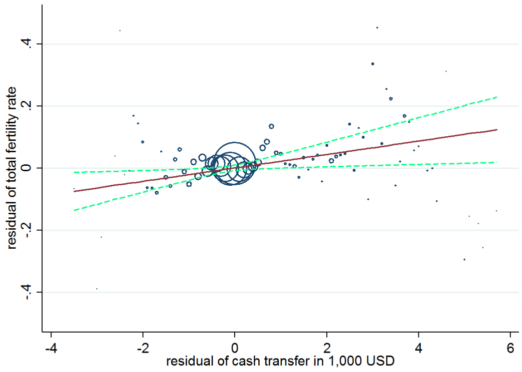


Figure 2.4: Placebo Test

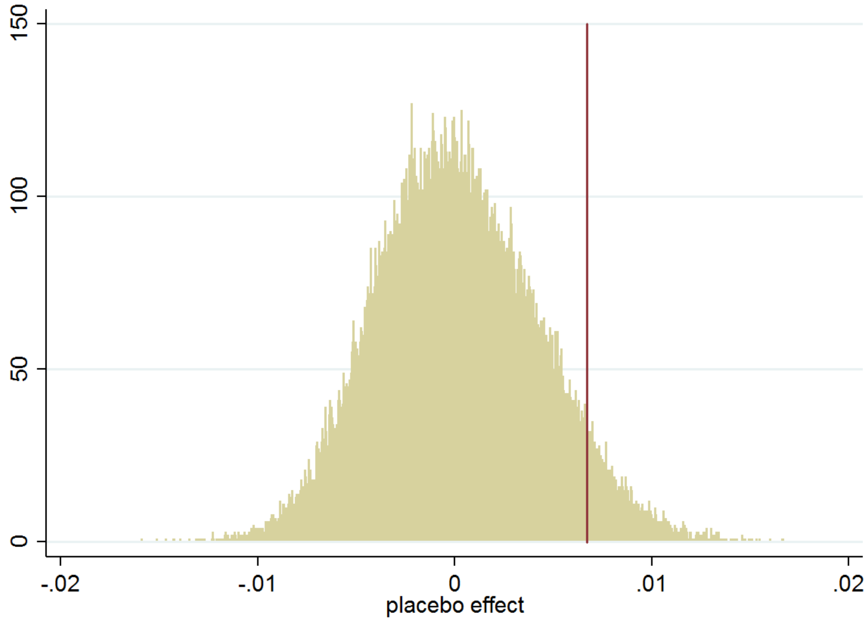
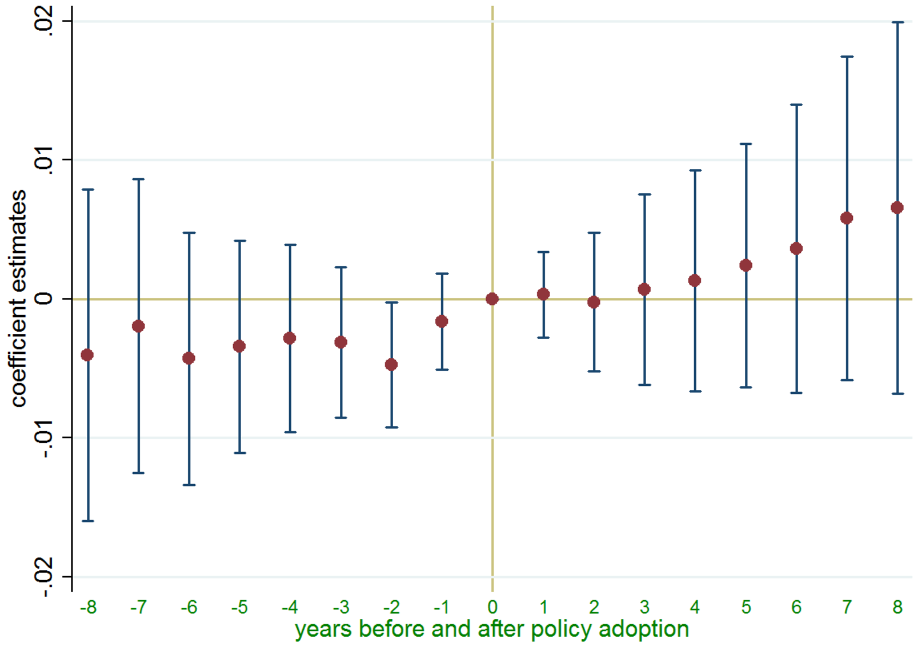


Figure 2.5: Health Measures at Birth Before and After Policy Implementation

(a) Birth Weight



(b) Pregnancy Duration

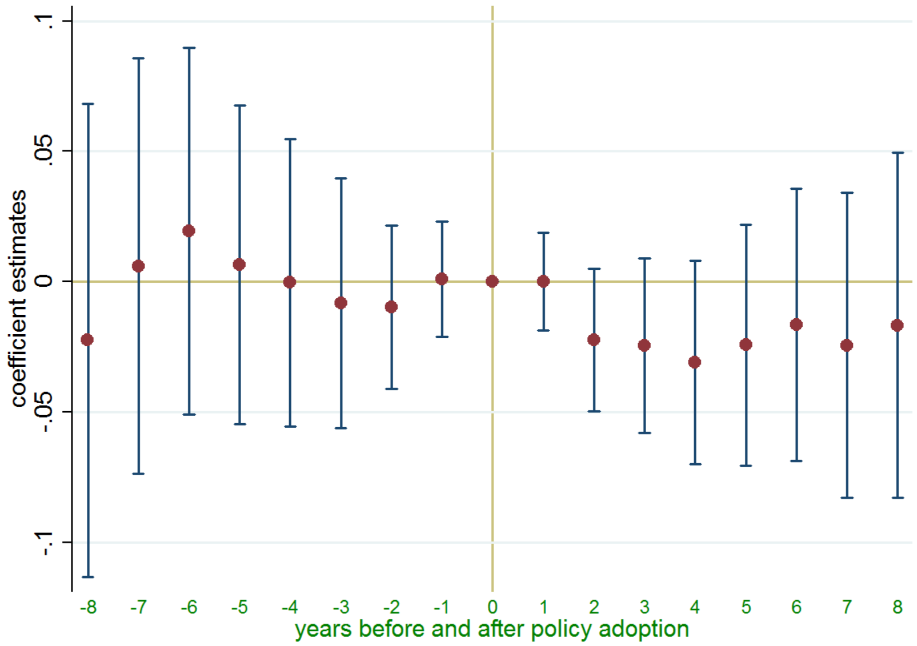
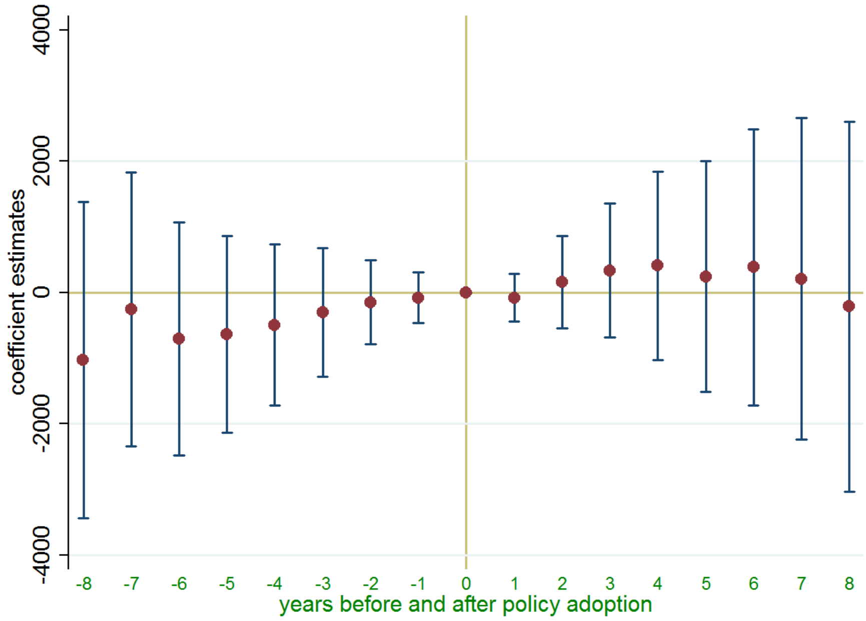
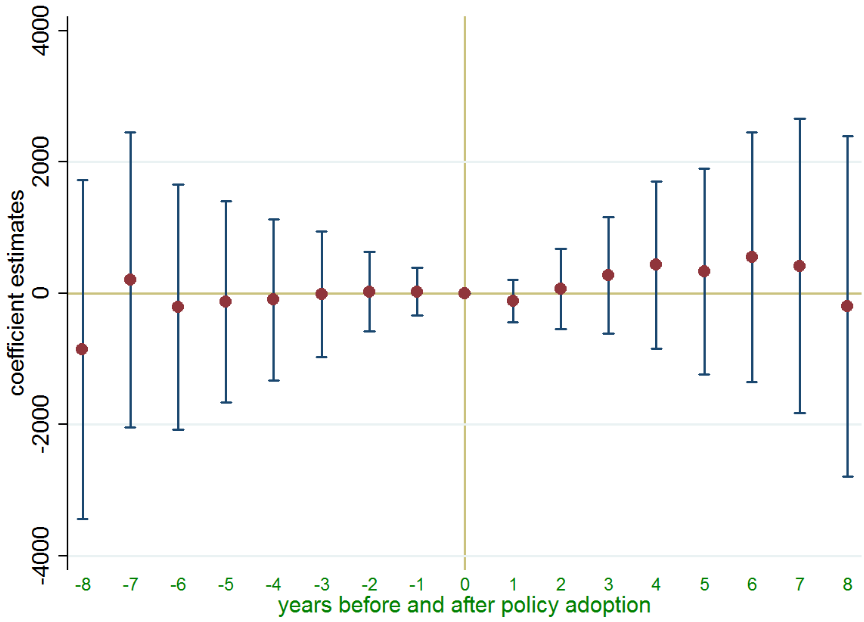


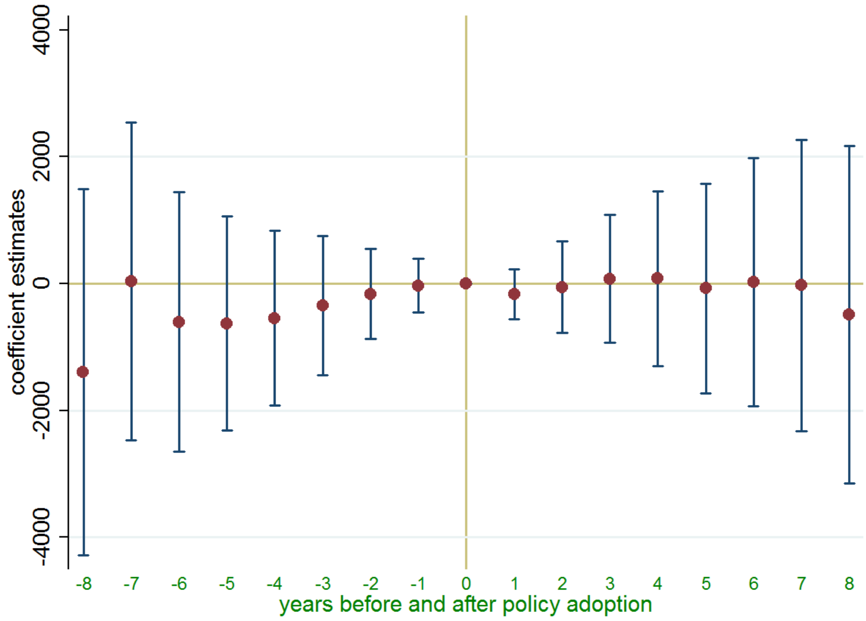
Figure 2.6: Robustness Checks

(a) Female Population between the ages of 15 and 19

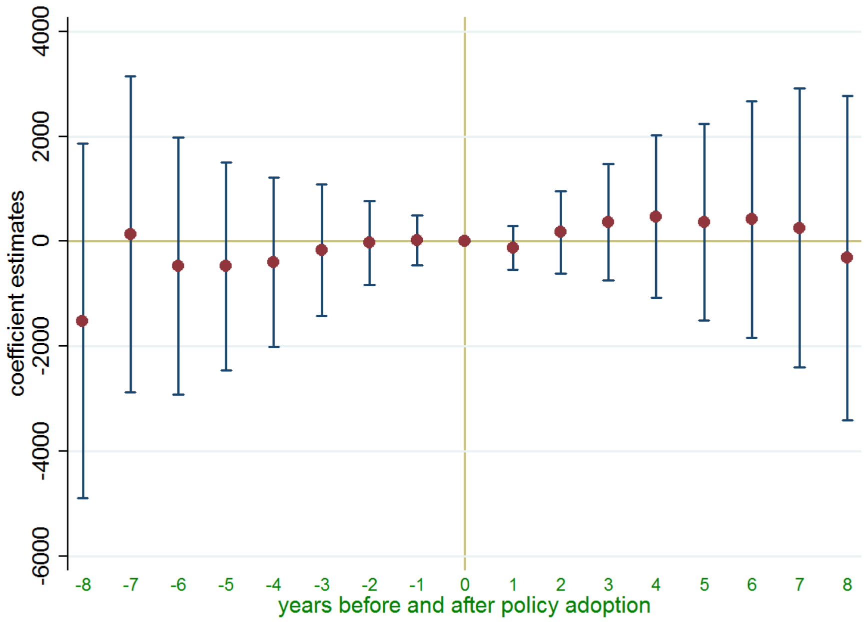


(b) Female Population between the ages of 20 and 24

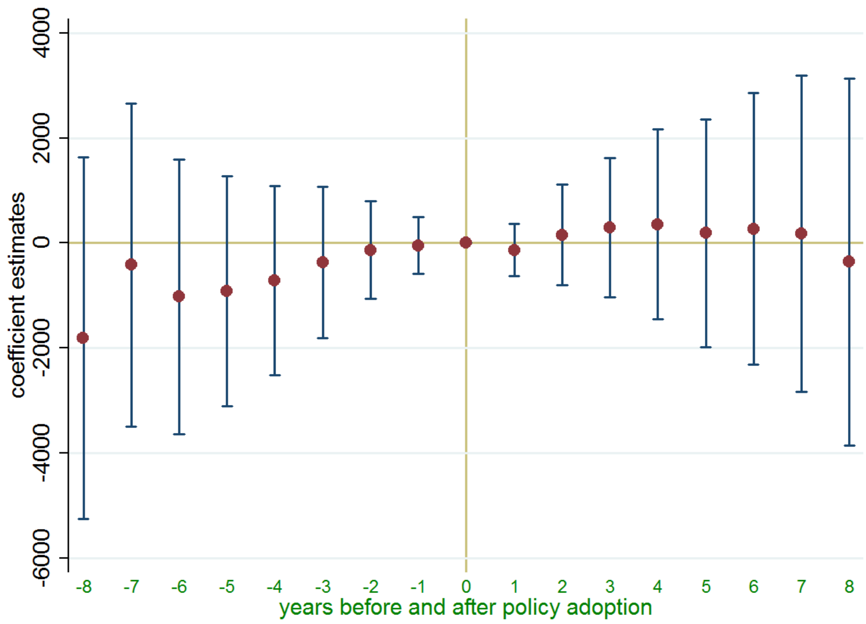




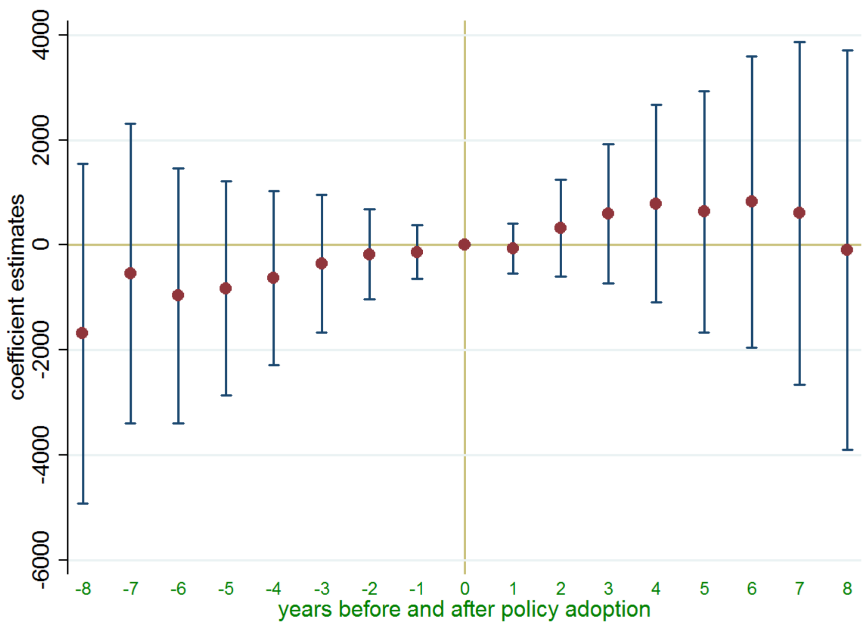
(c) Female Population between the ages of 25 and 29



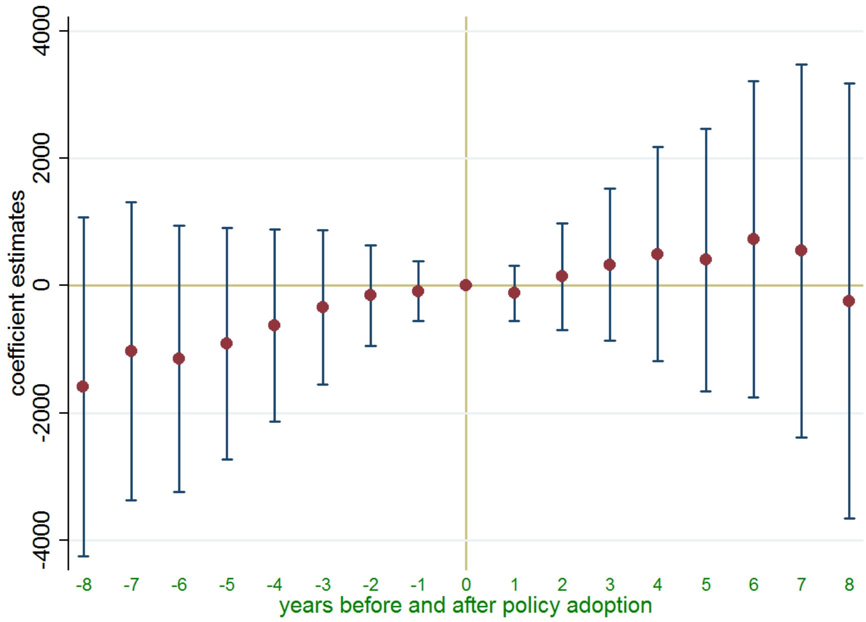
(d) Female Population between the ages of 30 and 34



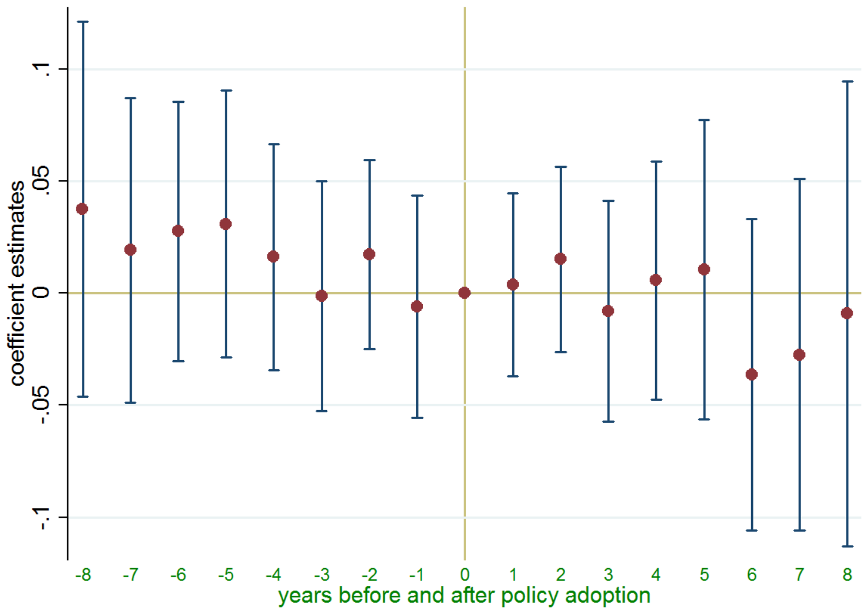
(e) Female Population between the ages of 35 and 39



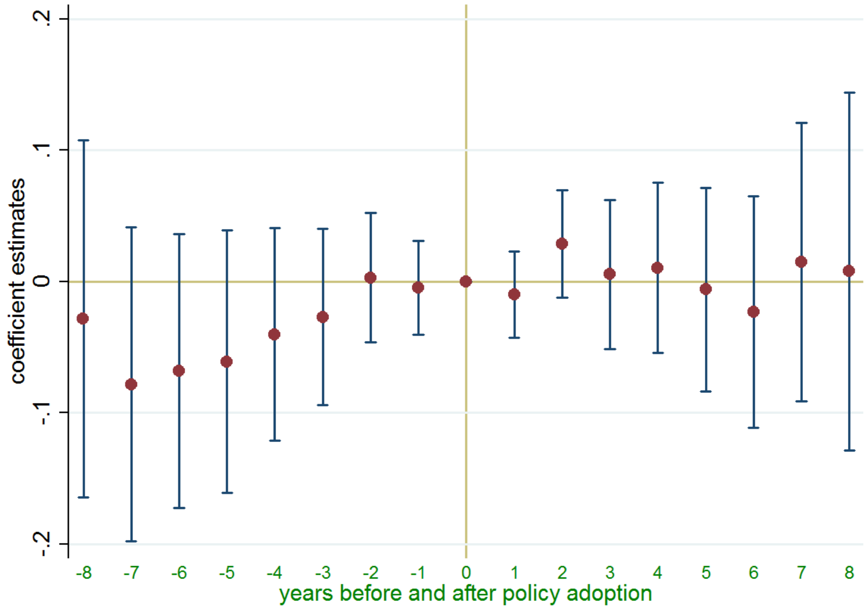
(f) Female Population between the ages of 40 and 44



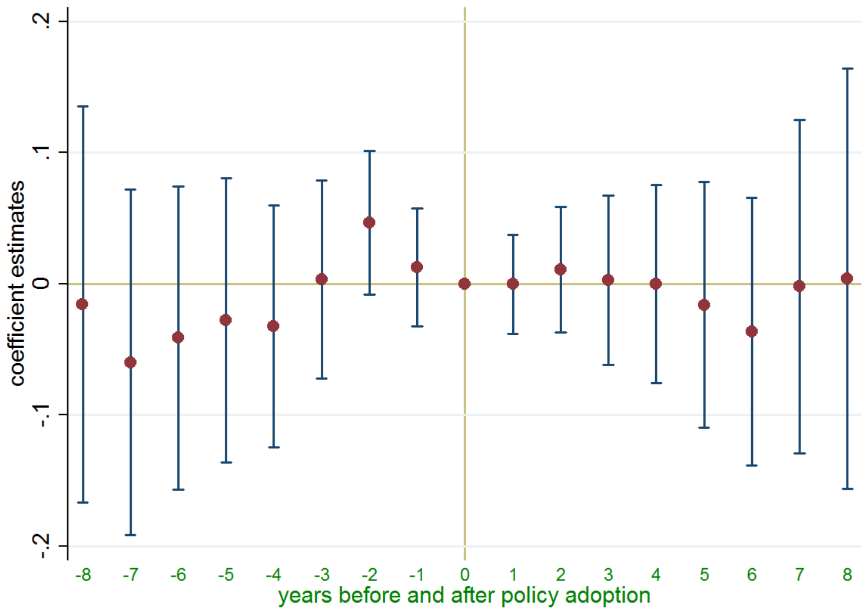
(g) Female Population between the ages of 45 and 49



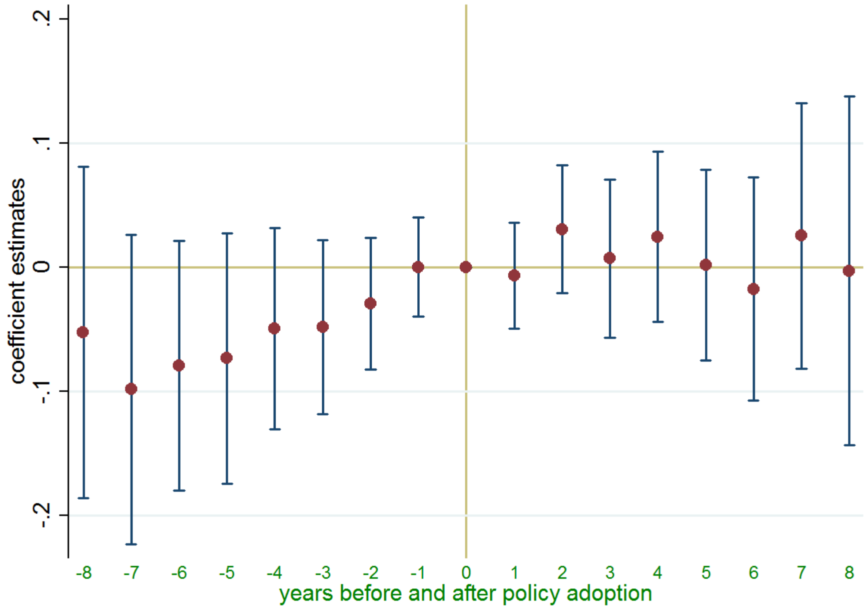
(h) Average Female Age at First Marriage



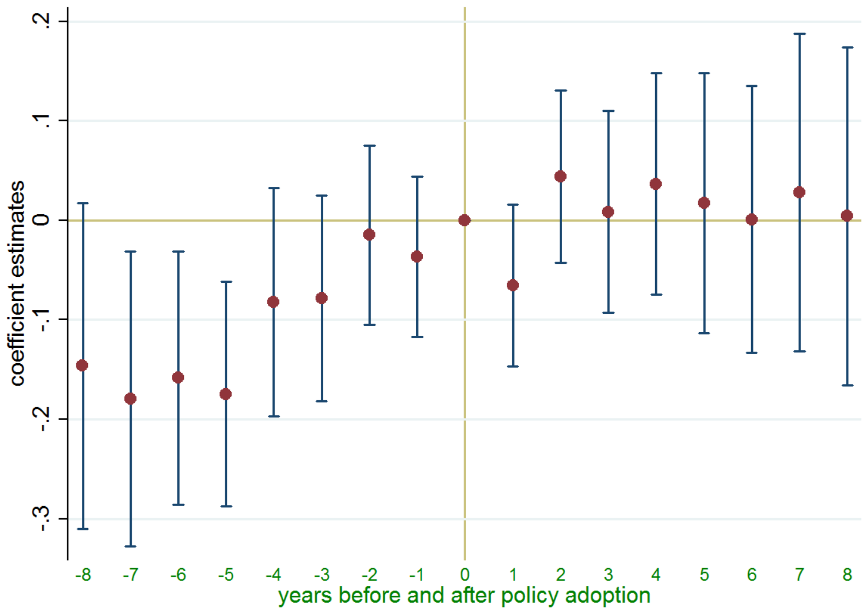
(i) Average Age of Mother



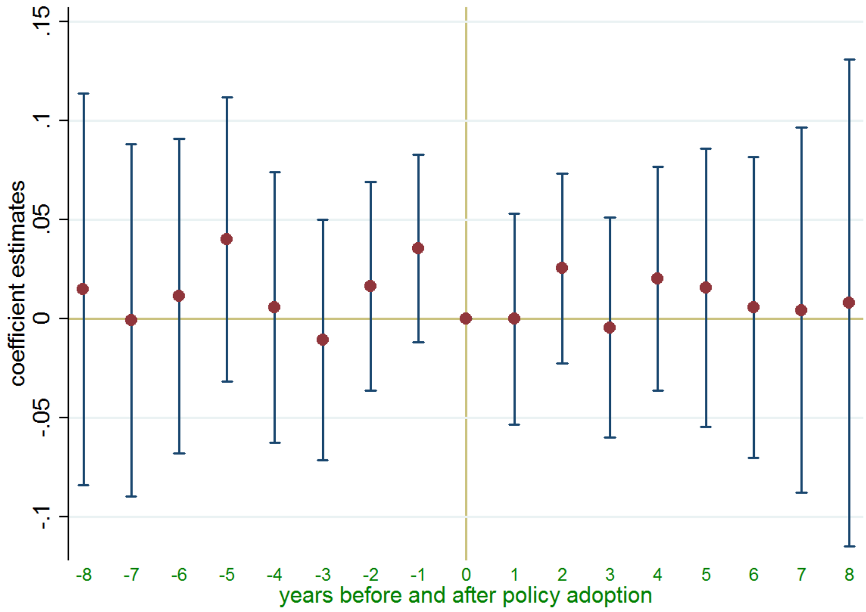
(j) Average Age of Mother at 1st Birth



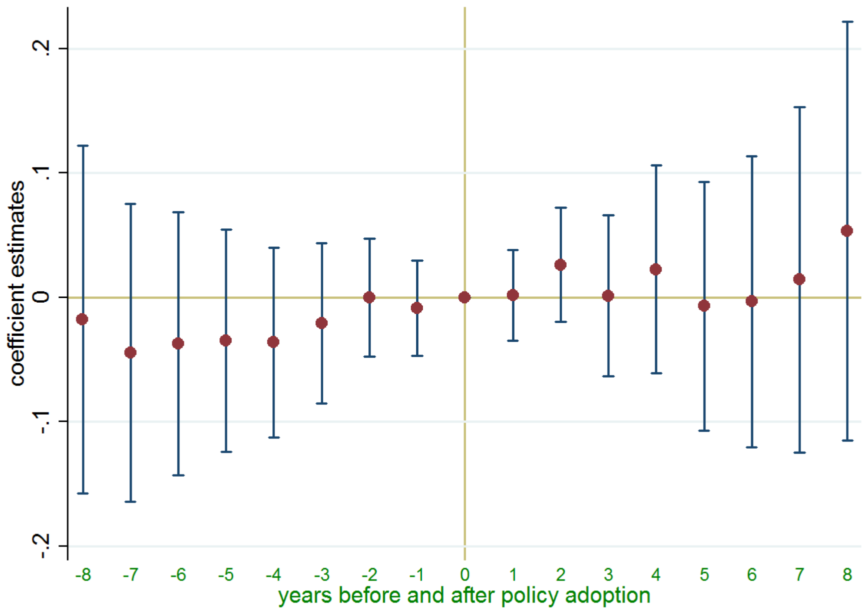
(k) Average Age of Mother at 2nd Birth



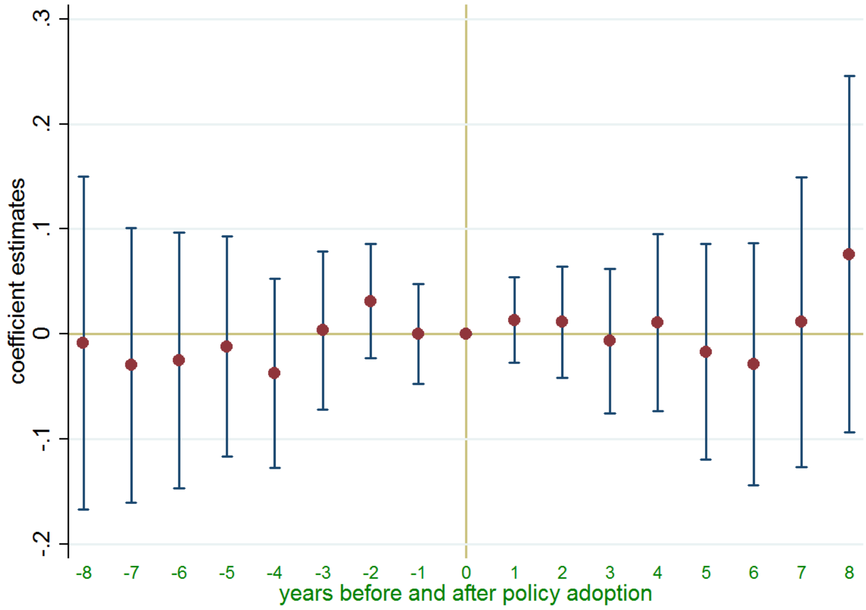
(l) Average Age of Mother at 3rd Birth



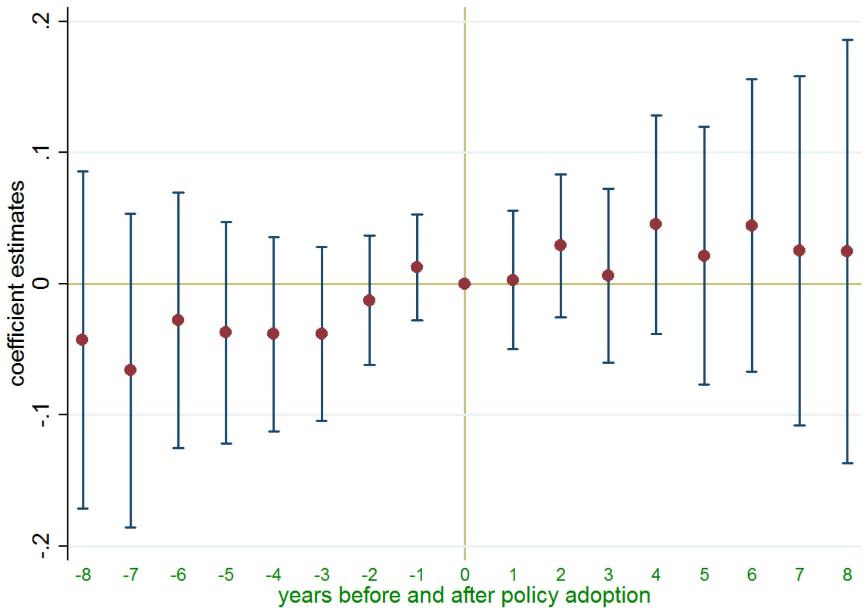
(m) Average Male Age at First Marriage



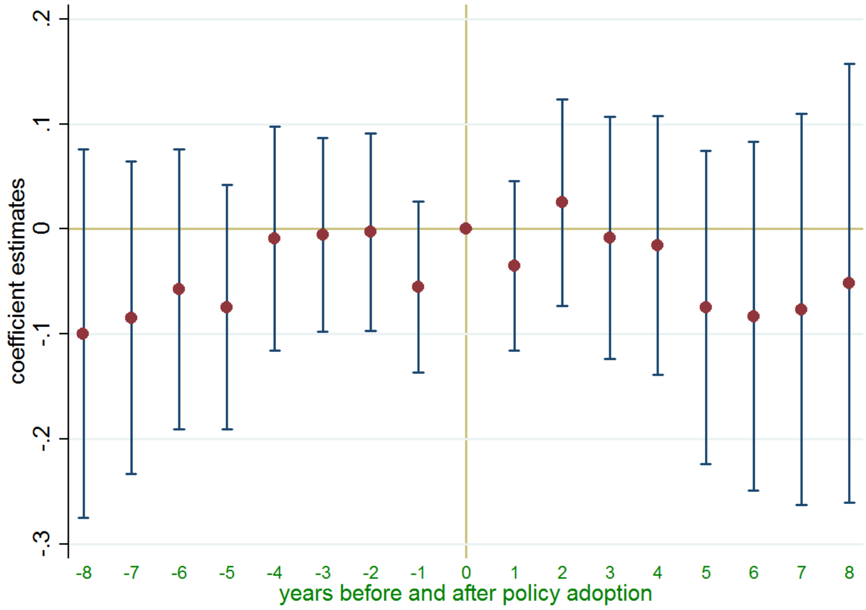
(n) Average Age of Father



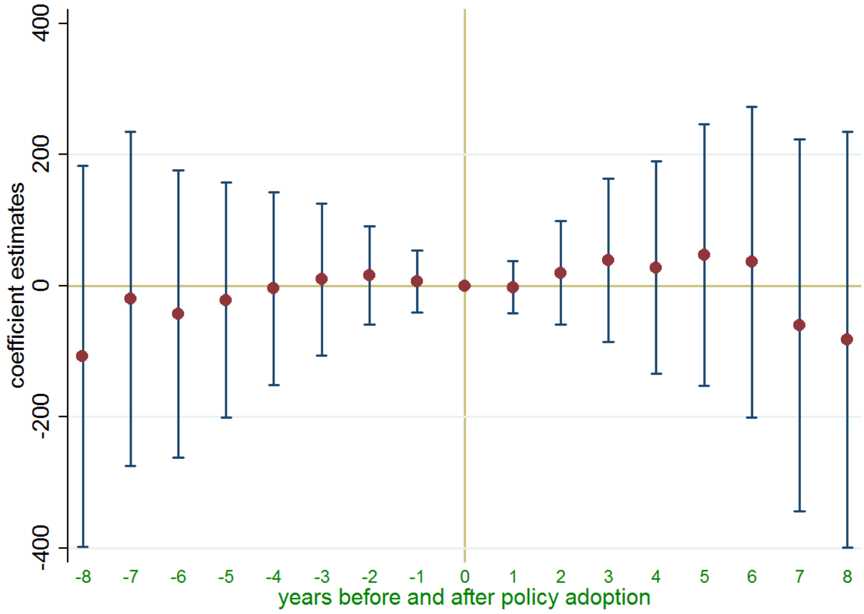
(o) Average Age of Father at 1st Birth



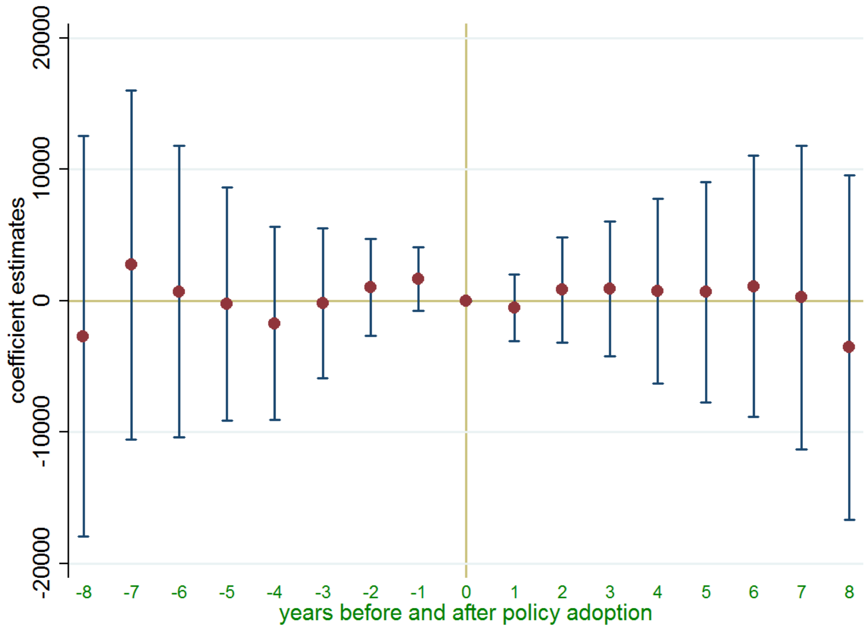
(p) Average Age of Father at 2nd Birth



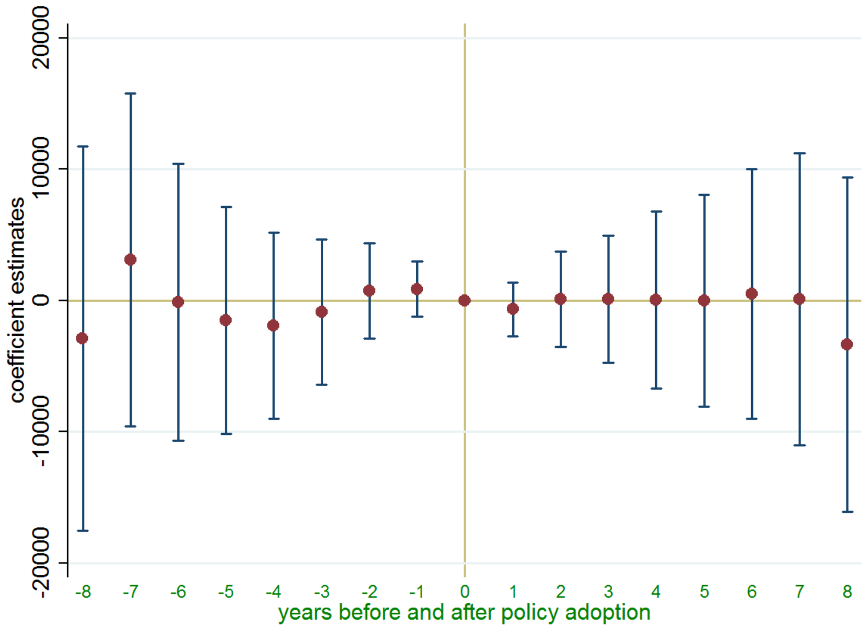
(q) Average Age of Father at 3rd Birth



(r) Number of Deaths



(s) Number of Move-Ins



(t) Number of Move-Outs

Table 2.13: Robustness Checks

	(1)	(2)	(3)	(4)
A. Dependent Variable:	Female Population of Age			
	Total	15-19	20-24	25-29
Cash Transfer ($CT_{d,p,t-1}$)	-574.5	-57.78	68.72	175.8
	(1,614)	(117.7)	(112.2)	(130.4)
Observations	3,230	3,230	3,230	3,230
R-squared	0.720	0.695	0.738	0.756
B. Dependent Variable:	Female Population of Age			
	30-34	35-39	40-44	45-49
Cash Transfer ($CT_{d,p,t-1}$)	83.88	-23.61	-79.84	-89.20
	(144.7)	(150.4)	(156.7)	(141.7)
Observations	3,230	3,230	3,230	3,230
R-squared	0.737	0.716	0.706	0.714
C. Dependent Variable:	Father's Age at			
	Any Birth	1st Birth	2nd Birth	3rd Birth
Cash Transfer for 1st Child		0.0232		
		(0.0636)		
Cash Transfer for 2nd Child			0.00973	
			(0.0376)	
Cash Transfer for 3rd Child	0.0189			-0.00350
	(0.0124)			(0.0180)
Observations	3,230	3,271	3,218	3,230
R-squared	0.957	0.960	0.950	0.750
D. Dependent Variable:	Male Age at			
	1st Marriage	# Deaths	# Move-Ins	# Move-Outs
Cash Transfer ($CT_{d,p,t-1}$)	-0.00114	-15.82	45.48	89.86
	(0.00639)	(14.95)	(576.8)	(560.0)
Observations	3,230	3,024	3,230	3,230
R-squared	0.960	0.610	0.727	0.740

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1; (ii) Across panels and columns, the same set of control variables and the province-by-year fixed effects are used as the fully specified model in Column 4 of Panel B of Table 2.1.

2.9 Appendix B: Determinants of Policy Implementation Timing and Cash Transfer Generosity

Many factors can be behind the background of a local government's decision to adopt the pro-natalist cash transfers. A local government may have been concerned with it history of low fertility rates and decided to adopt pro-natalist cash transfer policy. In this section, I formally study the determinants of the policy adoption and cash transfer generosity.

First, I consider a static model specified as below:

$$YR_{d,T} = \alpha_o + X'_{d,T}\beta + \psi_p + \nu_{d,T}, \tag{2.4}$$

for each year $T = 2000, ..., 2005$, where $YR_{d,T}$ denotes the number of years until district d adopts the transfer policy in year T. $X_{d,T}$ is a vector of district-level characteristics. I introduce the providence fixed effects ψ_p to account for provincial-level characteristics that may affect the dependent variable. Table 2.14 presents the results using year 2000 as a base year. Across the columns, I progressively add the fixed effects and explanatory variables. In Column (1), I begin with a set of demographic measures. Column (2) introduces the province fixed effects. I include district-government characteristics in Column (3) and other local characteristics such as local employment density and number of kindergartens in Column (4).[8] In Table 2.15, I repeat the same analysis for years until 2005. Note that the number of observations decreases across columns because I only consider the districts without the transfer program in each year.

After controlling for the providence fixed effects in Column (2) of Table 2.14, the coefficients on TFR become insignificant thereafter. The signs of the coefficients on population characteristics are fairly consistent, and the coefficients on percent adult population and death rate continue to increase and become significant in Column (5) along with area of land traded per 1000 people and financial independence rate. Holding everything else constant, as a local government is more financially independent, it is likely to adopt a pro-natalist cash transfer policy early.

The coefficients on area of land traded, number of firms and number of laborers per 1000 people are significant cross columns in Table 2.15. On the one hand, the first two are positively associated with revenue. Greater revenue generated from higher numbers of land transactions and local income may imply that a local government is less incentivized

[8] District-level characteristics correspond to observed local characteristics directly pertaining to local governments or factors that may influence local fiscal capacities.

to adopt pro-natalist cash transfer policy for more people and children. On the other hand, the coefficient on number of laborers per 1000 people are negative. Number of people working in a district is hardly equal to actual number of residents, and they are not pay local tax. Holding everything else kept constant, a local government may view greater number of laborers in its jurisdiction as a potential revenue source and want to attract them with pro-natalist policy.[9]

In sum, I note that some of the districts characteristics that are more related to government revenue and budget are more correlated with policy adoption decision. Total fertility rates do not seem to play a role in determining policy adoption timing.

Understanding the determinants of cash transfer amounts is challenging. I showed that policy adoption decision is closely related to potential local fiscal gains. However, a decision on the amount of cash transfer requires a local government to postulate associated costs as well as potential benefits. For instance, if its adult population is relatively large, expected cost of a pro-natalist cash transfer is large because there are high number of potential beneficiaries. In order to prevent budget deficit, a local government with high adult population is likely to set comparatively lower comparatively lower amounts. If financial dependence on the national government is high, there is a tension between the interest of increasing revenue via pro-natalist cash transfers and its budget constraint. If potential gains are greater than expected costs, a local government may set relatively higher cash transfer amounts. Taking the mechanism driven by revenue benefits, there may be strategic interactions among local governments within each province: competing for population. Note that it is unlikely for people to move across provinces just because of baby bonuses. In order to address this type of concern, I control for the province and year fixed effects.

I study the determinants of cash transfer amounts in two ways. I focus on samples with strictly greater than 0 cash transfer amounts. Under this truncation approach, I forgo the information during when districts do not have pro-natalist cash transfer policy. The dependent variable is cash transfer amounts in 1,000,000 KRW. The explanatory variables are lagged one year as cash transfer amounts are generally determined prior to implementation. Table 2.16 summarizes the results. Like in the case of policy adoption decision, total fertility rates do not seem to play a role in determining generosity. This is understandable again because the concerns about the low fertility rates are mutual across

[9] In addition, I consider a probit model to study the probability of policy adoption, defining the dependent variable as an indicator taking the value of one if a district has a cash transfer policy and using the same set of district characteristics. Taking bureaucratic inertia and administrative lag (e.g. enactment process can take up to a few months) into account, I lag the district characteristics by one year. The results are available upon request: they are similar to those reported in Table 2.14 and Table 2.15.

districts and district governments' decision is driven by revenue benefits via pro-natalist policy. As predicted, percentage of adult population lowers cash transfer amount.[10]

Unlike the case of policy adoption, the variables associated with revenue and budget benefits are no longer significant. There are two factors that are worth mentioning. The coefficients on conservative party is negative and significant. Next, the number of childcare facilities run by local governments per 1000 children reduce the cash transfer amount. If there are a greater number of childcare facilities, districts may be financially constrained to provide less cash transfers, or may find it only necessary to relatively lower cash transfers given its extant child benefits available in its jurisdiction.

[10] Table 2.17 reports the results of a similar analysis based on Heckman's control function approach (Heckman, 1976). The results closely resemble the estimates using the truncation method, reported in Table 2.16.

Table 2.14: Determinants of Policy Adoption I (Base Year: 2000)

Dependent Variable: Years till Adoption from 2000	(1)	(2)	(3)	(4)
Total Fertility Rate	-2.945***	-0.759	-0.667	-0.657
	(1.130)	(0.996)	(1.020)	(1.033)
% Female Population	-17.59	4.050	9.119	4.787
	(14.86)	(14.78)	(14.61)	(15.07)
% Adult Population	14.73	12.08	20.86**	17.24*
	(9.159)	(8.919)	(9.188)	(9.669)
Death Rate	-0.0804	-0.0520	-0.122*	-0.126*
	(0.0719)	(0.0608)	(0.0660)	(0.0731)
Marriage Rate	19.39	9.589	19.56	33.19
	(47.99)	(41.37)	(40.54)	(42.36)
Female Age at First Marriage	0.426	-0.562	-0.114	-0.191
	(0.394)	(0.357)	(0.371)	(0.377)
Male Age at First Marriage	-0.147	0.727*	0.768*	0.813*
	(0.444)	(0.422)	(0.432)	(0.435)
Log(Population Density)	-0.455***	-0.181	0.0207	-0.0852
	(0.144)	(0.171)	(0.183)	(0.202)
Net Migration Per 1000 People	-0.000440	0.00403	0.00734	0.00596
	(0.00731)	(0.00632)	(0.00630)	(0.00664)
Financial Independence Rate			-0.0359***	-0.0270*
			(0.0109)	(0.0143)
Per Capita Budget			0.0262	0.0655
			(0.184)	(0.226)
Conservative Party			0.339	0.397
			(0.271)	(0.279)
Land Trade per 1000 People			0.00414**	0.00402**
			(0.00178)	(0.00180)
# Firms per 1000 People				0.0122
				(0.00921)
# Laborers per 1000 People				-0.00248
				(0.00192)
# Kindergartens per 1000 Children				-0.448
				(0.542)
Observations	222	222	219	219
R-squared	0.260	0.526	0.565	0.570
Province FE	N	Y	Y	Y

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1; (ii) District characteristics in base year 2000 are used as covariates.

Table 2.15: Determinants of Policy Adoption I (Base Year: 2001-2006)

Dependent Variable:	(1)	(2)	(3)	(4)	(5)	(6)
Years till Adoption from year	2001	2002	2003	2004	2005	2006
Total Fertility Rate	-0.178	1.297	0.194	-1.702	1.053	-0.537
	(0.971)	(1.135)	(1.242)	(1.248)	(1.328)	(1.368)
% Female Population	0.0754	12.80	7.193	-9.428	13.02	1.005
	(14.60)	(14.58)	(16.65)	(15.84)	(16.37)	(15.28)
% Adult Population	14.75	15.68*	18.17*	14.71*	15.56*	6.966
	(9.226)	(8.960)	(9.573)	(8.494)	(8.921)	(8.941)
Death Rate	-0.0909	-0.0696	-0.0931	-0.0995	-0.0512	-0.173*
	(0.0815)	(0.0709)	(0.0758)	(0.0709)	(0.0902)	(0.103)
Marriage Rate	19.40	-7.629	5.001	21.45	-16.12	35.52
	(44.50)	(45.70)	(44.65)	(39.07)	(72.74)	(80.78)
Female Age at First Marriage	-0.653*	-0.198	-0.00814	0.439	0.326	-0.201
	(0.394)	(0.367)	(0.379)	(0.369)	(0.343)	(0.352)
Male Age at First Marriage	0.776*	0.00387	0.412	-0.295	-0.169	-0.123
	(0.399)	(0.307)	(0.401)	(0.321)	(0.272)	(0.330)
Log(Population Density)	-0.110	-0.0337	-0.0211	-0.239	-0.114	-0.200
	(0.201)	(0.201)	(0.217)	(0.212)	(0.205)	(0.184)
Net Migration per 1000 People	-0.0110	-0.0127*	0.00221	-0.00990	0.000540	0.00415
	(0.00783)	(0.00713)	(0.00865)	(0.00821)	(0.00932)	(0.00834)
Financial Independence Rate	-0.0184	-0.00602	-0.0312*	-0.0239	-0.0165	-0.00994
	(0.0149)	(0.0154)	(0.0180)	(0.0149)	(0.0141)	(0.0126)
Per Capita Budget	-0.0324	-0.0755	-0.0630	0.143	0.0537	0.0539
	(0.199)	(0.0886)	(0.0780)	(0.111)	(0.116)	(0.0762)
Conservative Party	0.379	0.920***	0.933***	0.818***	0.998***	-0.657*
	(0.275)	(0.311)	(0.331)	(0.311)	(0.300)	(0.395)
Female Governing Head		-0.146	0.790	1.124	1.102	0.247
		(1.115)	(0.952)	(1.081)	(0.991)	(1.358)
Land Trade per 1000 People	0.00339**	0.00362***	0.00463***	0.00352***	0.00306***	0.00395***
	(0.00137)	(0.00130)	(0.00125)	(0.000901)	(0.000799)	(0.000910)
# Firms per 1000 People	0.0142	0.0149*	0.0119	0.0161*	0.0130*	0.0142*
	(0.00879)	(0.00839)	(0.00934)	(0.00825)	(0.00763)	(0.00723)
# Laborers per 1000 People	-0.00268	-0.00286*	-0.00171	-0.00303*	-0.00247	-0.00288**
	(0.00178)	(0.00167)	(0.00193)	(0.00168)	(0.00155)	(0.00135)
# Kindergartens per 1000 Children	-0.276	-0.105	-0.0897	-0.454	-0.374	-0.517
	(0.519)	(0.426)	(0.454)	(0.393)	(0.439)	(0.417)
Observations	220	218	198	197	174	154
R-squared	0.574	0.590	0.309	0.360	0.342	0.466

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1; (ii) District characteristics in the corresponding base year are used as covariates.

Table 2.16: Determinants of Cash Transfer Generosity I

Dependent Variable: Cash Transfer ($CT_{d,p,t}$)	(1)	(2)	(3)	(4)
Total Fertility Rate	1.255*	0.847	0.714	0.775
	(0.747)	(0.895)	(0.900)	(0.940)
% Female Population	32.17**	19.44	16.20	15.75
	(14.09)	(13.95)	(13.83)	(13.83)
% Adult Population	-0.0756	-7.668	-11.41*	-11.47*
	(5.699)	(6.199)	(6.596)	(6.724)
Death Rate	0.0679	0.0708	0.0865	0.0884
	(0.0722)	(0.0483)	(0.0615)	(0.0627)
Marriage Rate	-28.94**	-9.893	-20.31	-18.80
	(12.90)	(12.57)	(14.38)	(21.13)
Female Age at First Marriage	0.516***	0.0338	-0.0280	-0.0530
	(0.188)	(0.187)	(0.189)	(0.201)
Male Age at First Marriage	0.192	0.374	0.307	0.312
	(0.234)	(0.273)	(0.264)	(0.264)
Log(Population Density)	-0.370***	-0.0677	-0.0316	-0.0772
	(0.136)	(0.165)	(0.172)	(0.188)
Net Migration per 1000 People	-0.00121	-0.00178	-0.00295	-0.00324
	(0.00593)	(0.00500)	(0.00503)	(0.00484)
Financial Independence Rate			0.0202**	0.0282**
			(0.00900)	(0.0123)
Per Capita Budget			0.0141	0.0179
			(0.0648)	(0.109)
Conservative Party			-0.421**	-0.420**
			(0.195)	(0.193)
Female Governing Head			0.0732	0.0242
			(0.401)	(0.390)
Land Trade per 1000 People			3.48e-05	8.23e-05
			(0.00122)	(0.00116)
# Firms per 1000 People				0.00687
				(0.00706)
# Laborers per 1000 People				-0.00129
				(0.00114)
# Kindergartens per 1000 Children				-0.0808
				(0.547)
Observations	1,987	1,987	1,986	1,986
R-squared	0.303	0.474	0.480	0.482
Province-by-Year FE	N	Y	Y	Y

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1; (ii) Observations one year prior to policy adoption and onwards are used; (iii) The covariates are lagged a year.

Table 2.17: Determinants of Cash Transfer Generosity II

Dependent Variable: Cash Transfer ($CT_{d,p,t}$)	(1)	(2)	(3)	(4)
Total Fertility Rate	1.448*	0.699	0.649	0.690
	(0.777)	(0.969)	(0.967)	(0.999)
% Female Population	31.72**	20.06	17.28	17.21
	(14.04)	(14.11)	(13.73)	(13.97)
% Adult Population	-2.135	-7.461	-10.98*	-11.13
	(5.540)	(6.203)	(6.497)	(6.792)
Death Rate	0.0523	0.0545	0.0841	0.0835
	(0.0730)	(0.0444)	(0.0588)	(0.0598)
Marriage Rate	-21.66	-8.903	-20.03	-17.96
	(13.30)	(13.13)	(15.24)	(20.47)
Female Age at First Marriage	0.654***	0.164	0.109	0.0946
	(0.181)	(0.219)	(0.226)	(0.242)
Male Age at First Marriage	0.298	0.126	0.0827	0.0870
	(0.254)	(0.281)	(0.277)	(0.278)
Log(Population Density)	-0.367***	-0.0408	-0.0173	-0.0269
	(0.135)	(0.149)	(0.156)	(0.170)
Net Migration per 1000 people	-0.000822	-0.00115	-0.00244	-0.00238
	(0.00594)	(0.00481)	(0.00475)	(0.00467)
Financial Independence Rate			0.0207**	0.0216*
			(0.00861)	(0.0111)
Per Capita Budget			-0.0161	-0.00199
			(0.0670)	(0.0968)
Conservative Party			-0.382**	-0.361*
			(0.188)	(0.190)
Female Governing Head			-0.0458	-0.208
			(0.432)	(0.468)
Land Trade per 1000 people			9.83e-05	0.000128
			(0.000863)	(0.000825)
# Firms per 1000 People				0.00290
				(0.00622)
# Laborers per 1000 People				-0.000348
				(0.00104)
# Kindergartens per 1000 Children				-0.0958
				(0.502)
Observations	1,987	1,321	1,321	1,321
R-squared	0.306	0.412	0.417	0.417
Province FE	N	Y	Y	Y
Year FE	N	Y	Y	Y

Notes: (i) Clustered standard errors at the district level in parentheses: *** p<0.01, ** p<0.05, * p<0.1; (ii) Observations one year prior to policy adoption and onwards are used; (iii) The covariates are lagged a year.

References

Adao, Rodrigo, Michal Kolesár, and Eduardo Morales, "Shift-Share Designs: Theory and Inference," *The Quarterly Journal of Economics*, Forthcoming, pp. 1–44.

Ahlfeldt, Gabriel M., Stephen J. Redding, Daniel M. Sturm, and Nikolaus Wolf, "The Economics of Density: Evidence from the Berlin Wall," *Econometrica*, 2015, *83* (6), 2127–2189.

Ákao Valentinyi and Berthold Herrendorf, "Measuring Factor Income Shares at the Sectoral Level," *Review of Economic Dynamics*, 2008, *11*, 820–835.

Akcigit, Ufuk, Salome Baslandze, and Stefanie Stantcheva, "Taxation and the International Mobility of Inventors," *American Economic Review*, 2016, *106* (10), 2930–2981.

Albert, Christoph and Joan Monras, "Immigration and Spatial Equilibrium: the Role of Expenditures in the Country of Origin," *Working Paper*, 2019, pp. 1–86.

Albouy, David, "The Unequal Geographic Burden of Federal Taxation," *Journal of Political Economy*, 2009, *117* (4), 635–667.

_ , "Evaluating the Efficiency and Equity of Federal Fiscal Equalization," *Journal of Public Economics*, 2012, *96*, 824–839.

Amarante, Verónica, Marco Monacorda, Edward Miguel, and Adrea Vigorito, "Do Cash Transfers Improve Birth Outcomes/ Evidence from Mathed Vital Statistics and Program and Social Security Data," *American Ecnomic Journal: Economic Policy*, 2016, *8*, 1–43.

Andersson, Gunnar and Ann-Zofie Duvander, "Gender Equality and Fertility in Sweden," *Marriage & Family Review*, 2006, *39*, 121–142.

Angrist, Joshua, Victor Lavy, and H Gregg Lewis, "Multiple Experiments for the Causal Link between the Quantity and Quality of Children," *Journal of Local Economics*, 2010, *28*, 773–823.

Balboni, Clare, "In Harm's Way? Infrastructure Investments and the Persistence of Coastal Cities," *Working Paper*, 2019, pp. 1–86.

Bartik, Timothy J., *Who Benefits from State and Local Economic Development Policies?*, Upjohn Press, W.E. Upjohn Institute for Employment Research, 1991.

Becker, Gary S and H Gregg Lewis, "On the Interaction between the Quantity and Quality of Children," *Journal of Political Economy*, 1973, *81* (2), 279–288.

_ **and Nigel Tomes**, "Child Endowments and the Quantity and Quality of Children," *Journal of Political Economy*, 1976, *84*, 143–162.

Bertrand, Marianne, Esther Duflo, and Sendhil Mullainathan, "How Much Should We Trust Differences-In-Differences Estimates?," *The Quarterly Journal of Economics*, February 2004, *119* (1), 249–275.

Bianchi, Nicola, Michela Giorcelli, and Enrica Maria Martino, "The Effects of Fiscal Decentralization on Publicly Provided Services and Labor Markets," *Working Paper*, 2019, pp. 1–40.

Black, Dan, Natalia Kolesnikova, Seth Sanders, and Lowell Taylor, "Are Children "normal"?," *Federal Reserve Board of St. Louis Working Paper 2008-040C*, 2008.

Black, Sandra E., "Do Better Schools Matter? Parental Valuation of Elementary Education," *The Quarterly Journal of Economics*, May 1999, *114* (2), 577–599.

Black, Sandra E, Paul J Devereux, and Kjell G Salvanes, "The More the Merrier? The Effect of Family Size and Birth Order on Children's Education," *The Quarterly Journal of Economics*, 2005, *120* (2).

Boccuzzo, Giovanna, Marcantonial Dalla Zuanna, Gianpiero Caltabiano, and Marzia Loghi, "The Impact of the Bonus at Bith on Reproductive Behaviour in a Lowest-Low Fertility Context: Friuli-Venezia Giula (Italy)," *Vienna Yearbook of Population Research*, 2008, *6*, 125–147.

Borusyak, Kirill, Peter Hull, and Xavier Jaravel, "Quasi-Experimental Shift-Share Research Designs," *NBER Working Paper No. 24997*, 2018, pp. 1–41.

Bryan, Gharad and Melanie Morten, "The Aggregate Productivity Effects of Internal Migration: Evidence from Indonesia," *Journal of Political Economy*, 2018, pp. 1–78.

Cadena, Brian and Brian Kovak, "Immigrants Equilibrate Local Labor Markets: Evidence from the Great Recession," *American Economic Journal: Applied Economics*, 2016, *8* (1), 257–290.

Caliendo, Lorenzo, Maximiliano Dvorkin, and Fernando Parro, "Trade and Labor Market Dynamics: General Equilibrium Analysis of the China Trade Shock," *Econometrica*, May 2019, *87* (3), 741–835.

Card, David, "Immigration Inflows, Native Outflows, and the Local Market Impacts of Higher Immigration," *Journal of Labor Economics*, 2001, *19* (1), 22–64.

Cellini, Stephanie R., Fernando Ferreira, and Jesse Rothstein, "The Value of School Facility Investments: Evidence from a Dynamic Regression Discontinuity Design," *The Quarterly Journal of Economics*, 2010, *125* (1), 215–261.

Chay, Kenneth Y. and Michael Greenstone, "Does Air Quality Matter? Evidence from the Housing Market," *Journal of Political Economy*, April 2005, *113* (2), 376–424.

Ciccone, Antonio and Robert Hall, "Productivity and the Density of Economic Activity," *American Economic Review*, 1996, *86* (1), 54–70.

Cohen, Alma, Rajeev Dehejia, and Dmitri Romanov, "Financial Incentives and FErtility," *The Review of Economics and Statistics*, 2013, *95* (1), 1–20.

Combes, Pierre and Laurent Gobillon, "The Empirics of Agglomeraion Economics," in Gilles Duranton, Vernon Henderson, and William Strange, eds., *Handbook of Regional and Urban Economics*, Amsterdam: Elsevier, 2015, chapter 5, pp. 247–348.

Congress of South Korea, "Local Tax Act (No.14033)," Online September 2016.

Crump, Richard, Gopi Shah Goda, and Kevin J Mumford, "Fertility and the Personal Exemption," *The American Economic Review*, 2011, *101* (4), 1616–1628.

Currie, Janet, *Poverty, the Distribution of income, and Public Policy*, New York: Russel Sage,

Davis, Marris A. and François Ortalo-Magné, "Household Expenditures, Wages, Rents," *Review of Economic Dynamics*, 2011, *14*, 248–261.

de la Roca, Jorge and Diego Puga, "Learning by Working in Big Cities," *Review of Economic Studies*, 2017, *84*, 106–142.

Desmet, Klaus and Esteban Rossi-Hansberg, "Urban Accounting and Welfare," *American Economic Review*, 2013, *103* (6), 2296–2327.

Diamond, Rebecca, "The determinants and welfare implications of US workers' diverging location choices by skill: 1980 to 2000," *American Economic Review*, 2016, *106* (3), 479–524.

Dix-Carneiro, Rafael, "Trade Liberalization and Labor Market Dynamics," *Econometrica*, May 2014, *82* (3), 825–885.

Donaldson, Dave and Richard Hornbeck, "Railroads and American Economic Growth: A Market Access Approach," *Quarterly Journal of Economics*, 2016, *131* (2), 799–858.

Ducruet, César, Réka Juhász, Dávid Krisztián Nagy, and Claudia Steinwender, "All Aboard: The Aggregate Effects of Port Development," *Working Paper*, 2019, pp. 1–70.

Eaton, Jonathan and Samuel Kortum, "Technology, Geography, and Trade," *Econometrica*, 2002, *70* (5), 1741–1779.

Edlund, Lena and Chulhee Lee, "Son Preference, Sex Selection and Economic Development: The Case of South Korea," *NBER Working Paper No. 18679*, 2013.

Eeckhout, Jan, Roberto Pinheiro, and Kurt Schmidheiny, "Spatial Sorting," *Journal of Political Economy*, 2014, *122* (3), 554–620.

Fajgelbaum, Pablo D. and Cecile Gaubert, "Optimal Spatial Policies, Geography and Sorting," *NBER Working Paper 24632*, 2018, pp. 1–64.

_ , Eduardo Morales, Juan Carlos Suarez-Serrato, and Owen Zidar, "State Taxes and Spatial Misallocation," *Review of Economic Studies*, 2019, *86* (1), 333–376.

Fisman, Raymond and Roberta Gatti, "Decentralization and Corruption: Evidence from Countries," *Journal of Public Economics*, 2002, *83*, 325–345.

Fleckenstein, Timo and Soohyn Christine Lee, "The Politics of Postindustrial Social Policy: Family Policy Reforms in Britain, Germany, South Korea, and Sweden," *Comparative Political Studies*, 2012, *47* (4), 601–630.

Frejka, Tomas, Gavin W Jones, and Jean-Paul Sardon, "East Asian Childbearing Patterns and Policy Developments," *Population and Development Review*, 2010, *36* (3), 579–606.

Gautheir, Anne, "The Impacts of Family Policies on Fertility in Industrialized Countries: A Review of the Literature," *Population Research Policy Review*, 2007, *26* (3), 323–346.

Gelbach, Jonah B., "Migration, the Life Cycle, and State Benefits: How Low Is the Bottom?," *Journal of Political Economy*, 2004, *112* (5), 1091–1130.

Glaeser, Edward and David Maré, "Cities and Skills," *Journal of Labor Economics*, 2001, *19* (2), 316–342.

_ and Joshua Gottlieb, "The Economics of Place-making Policies," *Brookings Papers on Economic Activity*, 2008, *Spring*, 155–239.

Goldsmith-Pinkham, Paul, Isaac Sorkin, and Henry Swift, "Bartik Instruments: What, When, Why, and How," *NBER Working Paper 24408*, 2018, pp. 1–66.

Gozález, Libertad, "The Effect of a Universal Child Benefit on Conceptions, Abortions, and Early Maternal Labor Supply," *American Economic Journal: Economic Policy*, 2013, *5* (3), 160–188.

Harper, Sarah, "Economic and Social Implications of Aging Societies," *Science*, 2014, *346* (6209), 587–591.

Heckman, James, "The Common Structure of Statistical Models of Truncations, Smaple Selection and Limited Development Variables and a Simple Estimator for Such Models," *Annals of Economic and Social Measurement*, 1976, *5*, 475–492.

Hong, Sok Chul, Young-Il Kim, Jae-Young Lim, and Mee-Young Yeo, "Pronatalist Cash Grants and Fertility: A Panel Analysis," *The Korean Economic Review*, 2016.

Kim, Doo-Rae, "Local Government Policy Diffusion in a Decentralised System," *Local Government Studies*, 2013, *32* (4), 582–599.

Kleven, Henrik and Wojciech Kopczuk, "Transfer Program Complexity and the Take-up of Social Benefits," *American Economic Journal: Economic Policy*, 2011, *3* (1), 54–90.

Kleven, Henrik Jacobsen, Camille Landais, Emmanuel Saez, and Esben Schultz, "Migration and Wage Effects of Taxing Top Earners: Evidence from the Foreigners' Tax Scheme in Denmark," *The Quarterly Journal of Economics*, 2014, pp. 333–378.

Kline, Patrick and Enrico Moretti, "People, Places, and Public Policy: Some Simple Welfare Economics of Local Economic Development Programs," *The Annual Review of Economics*, 2014, *6*, 629–662.

Lalive, Rafael and Josef Zweimüller, "How Does Parental Leave Affect Fertility and Return to Work?," *The Quarterly Journal of Economics*, 2009, *124* (3), 1363–1402.

Lee, Jungmin, "The Labor Market in South Korea, 2000-2016," Technical Report, IZA 2017.

Lee, Sam-Sik, "Low Fertility and Policy Responses in Korea," *the Japanese Journal of Population*, 2009, *7* (1), 57–70.

Liu, Haoming, "The Quality-Quantity Trade-Off: Evidence from the Relaxation of China's One-Child Policy," *Journal of Population Economics*, 2014, *27* (2), 565–602.

Manning, Alan and Barbara Petrongolo, "How Local Are Labor Markets? Evidence from a Spatial Job Search Model," *American Economic Review*, 2017, *107* (10), 2877–2907.

Marinescu, Ioana and Roland Rathelot, "Mismatch Unemployment and the Geography of Job Search," *American Economic Journal: Macroeconomics*, July 2018, *10* (3), 42–70.

McFadden, Daniel, "The Measurement of Urban Travel Demand," *Journal of Public Economics*, 1974, *3* (4), 303–328.

Milligan, Keven, "Subsidizing the Stork: New Evidence on Tax Incentives and Fertility," *The Review of Economics and Statistics*, 2005, *87* (3), 539–555.

Mogstad, Magne and Matthew Wiswall, "Testing the Quantity-Quality Model of Fertility: Estimation Using Unrestricted Family Size Models," *Quantitative Economics*, 2016, *7*, 157–192.

Molloy, Raven, Christopher L. Smith, and Abigail Wozniak, "Internal Migration in the United States," *Journal of Economic Perspective*, 2011, *25* (3), 173–196.

Monte, Ferdinando, Stephen Redding, and Esteban Rossi-Hansberg, "Commuting, Migration, and Local Employment Elasticities," *American Economic Review*, 2018, *108* (12), 2855–3890.

Moretti, Enrico and Daniel J. Wilson, "The Effect of State Taxes on the Geographical Location of Top Earners: Evidence from Star Scientists," *American Economic Review*, 2017, *107* (7), 1858–1903.

Morgan, Phillip S, "Is Low Fertility a Twenty-First-Century Demogrphic Crisis?," *Demography*, 2003, *40* (4), 589–603.

Morten, Melanie and Jacqueline Oliveira, "The Effects of Roads on Trade and Migration: Evidence from a Planned Capital City," *NBER Working Paper 22158*, 2018, pp. 1–64.

Moulton, Brent R., "An Illustration of a Pitfall in Estimating the Effects of Aggregate Variables on Micro Units," *The Review of Economics and Statistics*, May 1990, *72* (2), 334–338.

Newey, Whitney and Frank Windmeijer, "Generalized Method of moments with Many Weak Moment Conditions," *Econometrica*, 2009, *77* (3), 687–719.

Oates, Wallace E., "An Essay on Fiscal Federalism," *Journal of Economic Literature*, 1999, *37* (3), 1120–1149.

OECD, "Better Life Index: Korea," Technical Report 2016.

_ , "General Government Spending (Indicator)," Technical Report, OCED 2019.

Pellegrina, Heitor S. and Sebastian Sotelo, "Migration, Specialization, and Trade: Evidence from the Brazilian March to the West," *Working Paper*, June 2019, pp. 1–43.

Piyapromdee, Suphanit, "The Impact of Immigration on Wages, Internal Migration, and Welfare," *Working Paper*, 2017, pp. 1–83.

Preston, Samuel, Patrick Heuveline, and Michel Guillot, *Demography: Measuring and Modeling Population Processes*, Wiley-Blackwell, October 2000.

Rappaport, Jordan, "Moving to Nice Weather," *Regional Science and Urban Economics*, 2007, *37*, 375–398.

Redding, Stephen and Esteban Rossi-Hansberg, "Quantitative Spatial Economics," *Annual Review of Economics*, 2017, *9*, 21–58.

_ **and Matthew Turner**, "Transportation Costs and the Spatial Orgnization of Economic Activity," in Gilles Duranton, Vernon Henderson, and William Strange, eds., *Handbook of Regional and Urban Economics*, Amsterdam: Elsevier, 2015, chapter 20, pp. 1339–1398.

Roback, Jennifer, "Wages, Rents, and the Quality of Life," *Journal of Political Economy*, 1982, *90* (6), 1257–1278.

Rosen, Sherwin, "Wages-based Indexes of Urban Quality of Life," *Current Issues in Urban Economics*, 1979, pp. 34–55.

Sanderson, Eleanor and Frank Windmeijer, "A Weak Instrument F-Test in Linear IV Models with Multiple Endogenous Variables," *Journal of Econometrics*, 2016, *190*, 212–221.

Schafer, Andreas, "Regularities in Travel Demand: an International Perspective," *Journal of Transportation Statistics*, 2000, *3*, 1–31.

Stock, James, Jonathan Wright, and Motohiro Yogo, "A Survey of Weak Instruments and Weak Identification in Generalized Method of Moments," *Journal of Business & Economic Statistics*, 2002, *20* (4), 518–529.

Strulik, Holger and Sebastian Vollmer, "The Fertility Transition Around the World," *Journal of Population Economics*, 2015, *28*, 31–44.

Suárez-Serrato, Juan Carlos and Philippe Wingender, "Estimating the Incidence of Government Spending," *Working Paper*, 2014, pp. 1–69.

_ **and** _ , "Estimating Local Fiscal Multipliers," *Working Paper*, 2016, pp. 1–99.

Tiebout, Charles, "A Pure Theory of Local Expenditures," *Journal of Political Economy*, 1956, *64* (5), 416–424.

Tsivanidis, Nick, "Evaluating the Impact of Urban Transit Infrastructure: Evidence from Bogotá's TransMilenio," *Working Paper*, 2019, pp. 1–56.

Whittington, Leslie A, James Alm, and Elizabeth Peters, "Fertility and the Personal Exemption: Implicit Pronatalist Policy in the United States," *The American Economic Review*, 1990, *80* (3), 545–556.

Wildasin, David, "Locational Efficiency in a Federal System," *Regional Science and Urban Economics*, November 1980, *10*, 453–471.

Wildasin, David E., "Theoretical Analysis of Local Public Economics," *Handbook of Regional and Urban Economics*, 1987, *2*, 1131–1178.